Charles University in Prague

Faculty of Social Sciences
Institute of Economic Studies

MASTER THESIS

Asymmetric Monetary Transmission? Evidence from CEE Region

Author: **Amela Saric**

Supervisor: **Mgr. Filip Rozsypal, M.Sc.**

Academic Year: **2011/2012**

Declaration of Authorship

The author hereby declares that he compiled this thesis independently, using only the listed resources and literature.

The author grants to Charles University permission to reproduce and to distribute copies of this thesis document in whole or in part.

Prague, July 31, 2012

Signature

Acknowledgments

I am greatly indebted to my supervisor, Mr. Filip Rozsypal, whose insightful and patient guidance steered my effort in the right direction.

Further, I am much obliged to Mr. Tomáš Havránek for the data he provided and to Mr. Jaromír Baxa for additional methodological consultations. Finally, I wish to thank Mr. Georgios Georgiadis from Goethe University of Frankfurt for the Matlab code. The usual caveat applies.

Abstract

This thesis investigates monetary transmission asymmetries in CEE region. The first part addresses the role of credit growth in monetary transmission in the Czech Republic. Employing Logistic Smooth Transition Vector Autoregression model over the 1998:M1-2012:M3 period, we find that high credit growth dampens the effectiveness of monetary policy. No asymmetries in relative effects of contractionary and expansionary monetary policy shocks have been documented. In the second part, we apply the variation of Panel VAR to examine the role of financial structure in monetary transmission. The analysis is conducted on a sample of eight CEE states, encompassing the 1999:Q1-2009:Q4 period. Higher credit dependence is found to enhance the interest rate pass-through. However, cross-country asymmetries vanish when the credit dependence is interacted with the measure of banking sector competition. The ultimate role of financial structure in output and price fluctuations is indeterminable.

JEL Classification	F12, F21, F23, H25, H71, H87
Keywords	Monetary transmission asymmetries, LSTVAR, Credit growth, PCHVAR, Financial structure

Contents

List of Tables

List of Figures

Acronyms

BLUE Best Linear Unbiased Estimator

CEE Central and Eastern Europe

CEMAC Central African Economic and Monetary Community

FAVAR Factor-Augmented Vector Autoregression Model

GIRF Generalized Impulse Response Function

HP Hodrick-Prescott Filter

IFS International Financial Statistics

IRF Impulse Response Function

LM Lagrange Multiplier

LSTVAR Logistic Smooth Transition Vector Autoregression Model

NPL Non-performing loans

OLG Overlapping Generations Model

OLS Ordinary Least Squares

QFR Quarterly Financial Report of Manufacturing Firms

SVAR Structural Vector Autoregression Model

TAR Threshold Autoregression Model

TVAR Threshold Vector Autoregression Model

TVP VAR Time-varying Parameter Vector Autoregression Model

VAR Vector Autoregression Model

VECM Vector Error Correction Model

Master Thesis Proposal

Author	Amela Saric
Supervisor	Mgr. Filip Rozsypal, M.Sc.
Proposed topic	Asymmetric Monetary Transmission? Evidence from CEE Region

Topic characteristics Literature on financial frictions recognizes credit channel as the key source of asymmetries. Failure of Miller-Modigliani theorem indicates that the relative impact of monetary shocks varies based on the underlying credit conditions. To rephrase, a negative monetary shock is likely to depress the spending more if the credit is already restricted than if it is highly available. Understanding the transmission mechanism is particularly important today, as the credit market has not fully recovered from the financial crisis. Therefore, I will investigate the degree to which the responses of macroeconomic variables to monetary policy shocks differ depending on the credit market conditions.

Studies on monetary transmission further indicate that the economy's credit dependence may enhance the interest rate pass-through. As a determinant of microeconomic decisions, higher credit dependence ought to increase the leverage by monetary authorities. I will inspect whether this claim holds and under which conditions.

Hypotheses

1. Tight credit conditions, indicated by the level of threshold variable below its critical value, amplify the effects of monetary shocks.
2. In downturns, contractionary monetary shocks are more potent than the expansionary ones.
3. Higher credit dependence fosters the monetary transmission.

Methodology The impact of credit shocks on output and prices is usually assessed using Vector Autoregression (VAR) models. However, simple VAR is not designed to capture possible nonlinear dynamics between the variables. This problem can be alleviated by Threshold Vector Autoregression (TVAR) model, which assumes the existence of two regimes. Monetary shocks will be assessed through generalized impulse responses, which allow for endogenous regime changes throughout the duration of the response.

Outline

1. Introduction
2. Theoretical Background
3. Empirical Evidence
4. Methodology
5. Results
6. Conclusion

Core bibliography

1. Atanasova, C. (2003): "Credit Market Imperfections and Business Cycle Dynamics: A Nonlinear Approach." *Studies in Nonlinear Dynamics and Econometrics* **7(4)**: pp. 5.

2. Balke, N. (2001): "Credit and Economic Activity: Credit Regimes and Non- linear Propagation of Shocks." *Review of Economics and Statistics* **7(4)**: pp. 344 – 349.

3. Bernanke, B. & M. Gertler (1989): "Agency Costs, Net Worth and Business Fluctuations." *American Economic Review* **79(2)**: pp. 14–31.

4. Calza, A. & J. Sousa (2005): "Output and Inflation Responses to Credit Shocks: Are there Threshold Effects in the Euro-area?" *Working Paper 481*, European Central Bank

5. Égert, B. & R. MacDonald (2008): "Monetary Transmission Mechanism in Central and Eastern Europe: Surveying the Surveyable" *OECD Economics Department Working Papers 654*, OECD Publishing

6. Weise, C.L. (1999): "The asymmetric effects of monetary policy: A nonlinear vector autoregression approach." *Journal of Money, Credit and Banking* **31(1)**: pp. 85–108.

Author | Supervisor

Chapter 1

Introduction

This essay investigates whether the monetary policy in CEE region may have asymmetric effects. Specifically, we evaluate the impact of monetary actions based on the underlying credit conditions and financial structure. Over the past two decades, CEE economies have been affected by numerous structural changes, including the transition to a market-based economy, trade integration with the European Union and global financial crisis. These developments have been complemented by the reforms aimed at promoting macroeconomic and financial stability. As an outcome, monetary transmission mechanism may have undergone certain changes. Our focus is on three key issues: propagation of monetary shocks at different stages of credit cycle, relative potency of contractionary versus expansionary monetary shocks and the cumulative impact of various components of financial structure on monetary transmission.

The first part of our analysis explores the relative potency of monetary shocks in the Czech Republic based on the credit market liquidity. Several papers provide evidence of asymmetries in specific aspects of monetary propagation, including the exchange rate channel and monetary policy rule (Babetskaia-Kukharchuk 2007; Horváth 2008b; Fidrmuc *et al.* 2008; Vašíček 2011). Our analysis is pursued within the framework of Logistic Smooth Transition Vector Autoregression (LSTVAR) model, which allows for asymmetric propagation of monetary policy shocks and endogenous regime switches without *a priori* assumptions about their particular dates. We track the evolution of monetary transmission mechanism over the longer sample period and with the methodology previously unapplied on CEE sample, thereby contributing to the existing empirical literature.

In the context of CEE region, one-country samples may be too short to provide robust conclusions. As advised by Jarociński (2008), we pool the data for eight CEE countries – the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Slovakia and Slovenia to test for the role of financial structure in monetary transmission. We employ Panel Conditionally Homogenous Vector Autoregression (PCHVAR) model over the 1999:Q1-2009:Q4 sample. The two proxies for financial structure are the economy's credit dependence and banking sector competition.

The rest of the essay is structured as follows. Chapter 2 surveys the most relevant economic theories behind the transmission channels. Chapter 3 pursues the LSTVAR analysis for the Czech Republic. The importance of financial structure for monetary transmission is tested by PCHVAR in the Chapter 4. Finally, Chapter 5 concludes.

Chapter 2

Remarks on Monetary Transmission Asymmetries

Surveying the process of monetary transmission is a daunting challenge. Not so long ago, Bernanke & Gertler (1995) attributed it a notion of the black box. The past two decades have seen a proliferation of empirical papers attempting to decipher the issue. Recent advances notwithstanding, this domain continues to preserve its original research appeal. A number of questions are yet to be answered. This issue comes to the fore in CEE region, where the much-needed research is still relatively sparse. Evaluating the power of monetary policy instrument to affect inflation or exchange rate is at the heart of all monetary arrangements. In the context of convergence with the euro area, understanding fundamental forces behind monetary transmission is of an utmost importance.

The notion of monetary policy asymmetries has been well established in macroeconomics (Blinder 1987; Bernanke & Gertler 1989; Bernanke *et al.* 1996). This chapter identifies two channels of monetary transmission which may be the source of asymmetries. We then turn our focus to the empirical evidence provided up-to-date.

2.1 Theoretical Background

2.1.1 Interest Rate Channel

Interest rate channel is the oldest and most conventional transmission channel. The propagation of monetary shocks is relatively simple: increase in policy rate entails higher deposit and loan rates, yields on government bills and bonds, along with the drop in equity prices. Higher nominal rate and price stickiness act to induce higher real rate, depressing investment, consumption and output.

A voluminous strand of literature identifies the sources of monetary transmission asymmetries in the mechanism outlined above. Ball & Mankiw (1994) and Tsiddon (1993) introduce the model with downward price rigidity. Assuming a change in monetary policy stance and the corresponding adjustment in trend inflation, firm's desired price level is altered. However, downward and upward adjustment entail different costs. Inflation depresses the firm's relative price, which renders the cost of downward adjustment lower. Hence, expansionary and contractionary monetary policy may have asymmetric effects on prices. Kandil (1995) applies the same line of reasoning to downward wage rigidity, citing its positive correlation with the price rigidity. Both studies conclude that nominal rigidities exacerbate the negative effects of contractionary demand shock on output.

Certain works ascribe monetary transmission asymmetries to the inflation rate (Shen & Chiang 1999; Mandler 2010). Monetary expansion stimulates the investment by lowering the interest rate, whereas higher level of investment acts to induce inflation. Consequently, increased inflationary expectations may invalidate the effects of monetary policy.

To analyze the degree of interest rate pass-through, it is instructive to decompose the transmission process into two distinct stages: (1) transmission from policy rate to long-term money market rate, and further to lending and deposit rates, (2) transmission to the real economy. The first stage crucially hinges upon the yield curve.[1] Transmission of interest rate from market to retail rates is described by marginal cost pricing model by Rousseas (1985):

$$i^B = \mu + \beta \cdot i^M, \tag{2.1}$$

where i^B is the bank lending/deposit rate, i^M represents the money market rate and μ stands for a constant markup. Changes in money market rate are fully propagated to the retail rates only if β equals 1. In the case of imperfect competition or information asymmetries, β will be different than unity.

The value of β depends on a range of factors. First, low elasticity of demand for loanable funds or deposits renders banks less responsive to market rate changes. This is pronounced in the absence of alternative sources of fi-

[1] Égert *et al.* (2007) cite four key determinants of the yield curve: (1) liquidity preference: less liquid long-term investment induce higher long-term interest rate, (2) market segmentation: short- and long-term rates may be determined at separate markets, (3) inflation expectations, (4) exchange rate expectations.

nancing or investment opportunities. Next, weak competitive pressures may insulate the retail rates from money market fluctuations, while more strenuous competition drives down the lending rate. The presence of asymmetric information may have two contrasting outcomes: credit rationing and higher interest rate. Further, the adjustment of lending rates is sluggish if there is no maturity mismatch.

Following the reaction to a monetary policy shock, interest rates impinge on the real economic activity. However, it is often argued that the interest rate channel alone cannot explain large shifts in economic activity emanating from monetary actions. This failure may emanate from flawed assumptions, namely that bonds and loans are perfect substitutes and that banks assume only a passive role of attracting the deposits. Moreover, it is important to disentangle between the shifts in loan demand and loan supply. Thus, we examine the functioning of credit channel.

2.1.2 Credit Channel

Credit channel rests on the assumption of financial frictions. In this respect, credit rationing model is instructive to begin with. Its crucial assumption is the information asymmetry between lenders and borrowers, who know little about each other. Banks play an intermediation role. Borrowing is collateralized due to an imperfect enforcement of loan contracts, whereas the value of collateral depends on asset prices. Shifts in expectations induce asset price fluctuations, shocks to the financial sector and eventually, their propagation to the real economy.

Bank Lending Channel

Bank lending channel propagates monetary policy shocks by affecting the amount of credit extended to firms and consumers. In one of the pioneering works in this domain, Stiglitz & Weiss (1981) coined the theory of credit rationing, which carries a strong message: loan market equilibrium may be characterized by credit rationing. To restate, if the equilibrium requires higher interest rate, banks may find it more profitable to restrict the credit instead. The reason is twofold. First, interest rate hikes depress the returns on successful projects, lowering banks' profits. Second, higher interest rate leads to an adverse selection problem: the prevalence of bad borrowers exacerbates the risk to the banks' portfolio. Thus, credit rationing appears as a more reasonable solution.

Monetary contraction shrinks the banks' balance sheets through open market operations or higher minimum reserve requirements, which consequently drain deposits from the system (Bernanke & Blinder 1988). Since the deposits are an essential source of loanable funds, monetary tightening induces lending cuts. This channel is operative under three conditions: banks are unable to insulate their portfolios from monetary tightening; consumers and firms are unable to fully substitute for the shortage of loans by raising the external sources of financing and price adjustment mechanism is imperfect. In a broader sense, monetary policy changes affect the cost of external funding, compromising the ability of banks and firms to maintain their loan portfolio.

A large body of evidence suggests that monetary policy impulses have distributional effects across the banks (Kashyap *et al.* 1993; Kashyap & Stein 1994; Kishan & Opiela 2000). First, small banks encounter more information asymmetries and find it more difficult to raise uninsured funds in the wake of monetary tightening. Thus, their lending cuts are disproportionately higher in comparison to the bigger banks. Second, monetary contraction erodes the bank capital, entailing a sharp lending contraction if the institution is highly leveraged. Finally, less liquid banks are more adversely affected, due to the lesser funds to counteract the negative monetary shocks. The outlined concepts hint that the economy's dependence on credit (i.e. the absence of alternative sources of financing) ought to result in a more complete interest rate pass-through.

Balance Sheet Channel

The idea of credit rationing was placed into the framework of financial accelerator mechanism. This concept relates to the notion that shocks to the real economy are amplified by deteriorating financial conditions. More precisely, higher cost of lending in downturns leads to an increase in uncertainty of projects and external finance premium. Consequently, lenders require higher collateral as a buffer against possible losses. External finance premium, which is inversely related to the borrower's net worth, is on a rise. Hence, borrowers are unable to provide the collateral required. In response, banks restrict their credit supply, which leads to production cuts. In good times, the opposite forces are at work: borrower's profits and net worth rise, agency costs decline and banks become more willing to provide funding.

Bernanke *et al.* (1996) argue that the firm's balance sheet serves both as a propagator and an amplifier of shocks. The effects of initial shock on production and spending are proportional to the change in borrowers' net worth. Thus, the propagation of good times ensues as a result of increasing net worth and abundant credit. On the contrary, shrinking balance sheets restrict the credit availability and reinforce the downturns.

In the studies by Kiyotaki & Moore (1997), Bernanke *et al.* (1998) and Moody & Taylor (2003), the idea of financial accelerator was expanded and modelled in a more sophisticated manner. Bernanke *et al.* (1998) build the dynamic general equilibrium model with financial frictions and four types of agents: households, entrepreneurs, retailers and government. Households lend money to the entrepreneurs, who sell their products to the retailers. Government conducts the monetary and fiscal policy. Intuitively, if Miller-Modigliani theorem does not hold, higher agency costs will be compensated by a higher external finance premium. On the other hand, external finance premium is highly countercyclical, due to the procyclicality of net worth. As an implication, downturns enhance swings in borrowing, which depresses investment, production and overall level of consumption. Hence, monetary policy shock is expected to exhibit a stronger impact on spending of credit-constrained firms. A general conclusion is that financial frictions magnify the effects of monetary shocks on investment and output.

This concept is widely applied in a number of papers. For instance, Kiyotaki & Moore (1997) describe the interaction of credit constraints and aggregate economic activity over the phases of business cycle. They develop a simple stylized equilibrium model, in which durable assets have a dual role, as a factor of production and a collateral for loans. Borrower's credit limits are affected by the asset prices. On the other hand, asset prices affect the borrower's net worth, the value of collateral and access to credit. Consequently, the most adversely affected agents are those who experience the highest decline in collateral prices. Holmstrom & Tirole (1997) introduce the model with limited borrowing capacity of both firms and banks. Funding sources are dependent on the firm's net worth. Due to higher agency costs, credit crunch hits highly leveraged firms more than the well-capitalized ones. This model is motivated by the financial crises underwent by several OECD economies throughout the 1980s and 1990s. Its predictions are consistent with the fluctuations observed in the examined period. Azariadis & Smith (1998) explore the transmission of credit shocks in traditional overlapping generations (OLG) setting. The economy lies either in a traditional Walrasian regime, characterized by unfettered operation of the credit markets, or the regime with binding credit constraints.

Regime switches occur as a result of changing market expectations: more pessimistic mood prompts the depositors to transfer their funds out of the banking sector for less productive purposes. Consequently, credit supply falls, shifting the equilibrium to a credit-constrained regime. An uplift in market expectations moves the economy back to the Walrasian regime. The study argues that the result is a set of equilibrium interest rates on loans that validate the depositors' original beliefs. Finally, Cooper & Corbae (2001) analyze financial collapses and their real implications through the prism of disrupted financial intermediation. Intermediary is viewed as a coalition of lenders, whose members share the fixed cost of lending. If a single agent believes that many others will join the coalition, his cost of lending declines, which in turn induces a high degree of participation. In glooms, the number of lenders decreases, reinforcing credit supply shortages. Thus, the intermediation is belief-driven: its success will depend on the confidence placed into the system. This model is able to replicate some of the key empirical facts of the Great Depression.

2.2 Evidence on the Interest Rate Channel

This section attempts to gather information on the determinants of pass-through from policy rate to money market rate, retail rates and to the real economy. This interplay has been explored in the context of financial structure, which is inseparable from the functioning of broad credit channel. Hence, we attempt to provide a comprehensive literature review below.

Many observers have noted that the retail interest rate exhibits downward rigidity. Specifically, market pressure induces a quick upward adjustment of the lending rate, while the downward adjustment is more sluggish. Similar relationship holds for borrowing rates. Weth (2002) finds that German banks were largely characterized by this behavior during the 1990s. However, the speed of interest rate adjustment varied depending on the banks' structural characteristics. Low market rates induce a drain on deposits. In a given situation, large banks are able to refinance quickly, due to the better capital market access. Hence, their interest rate is less downward rigid. Consequently, banks financed mainly by the savings deposits are less responsive to changes in the money market rate.

De Bondt (2005) tackles the issue more in-depth by considering the term structure of the interest rates. Euro area policy rate shifts are fully propagated to the short-term money market rates (up to three months). Monetary authorities' leverage diminishes with an increasing maturity of money market instrument. The pass-through from money market to retail rates is partial in the short-run, but substantial in the long-run. However, 100% propagation is detected only in the case of bank lending rates. The stickiest are the overnight and short-term deposit rates, with the long-term propagation amounting to 40% only. In a similar manner, Kwapil & Scharler (2006) draw a comparison of the interest rate pass-through between US and euro area. In both regions, the transmission of money market rate to retail rates is low for instruments with short maturities, but increases substantially with the term structure. Contrary to the household interest rates, corporate lending rates are heavily affected by the monetary policy shocks. Finally, the pass-through is found to be stronger in US than in the euro area.

Structure of the banking market and its ramifications for monetary transmission mechanism have been widely explored in the literature. The cornerstone study by Mojon (2000) analyzes six largest EU economies over the 1979-1998 time span, concluding that financial market deregulation and increased competition bolster the interest rate pass-through.[2] Further, it is found that higher volatility of money market rates impinges on monetary transmission.

Gropp *et al.* (2007) study the pass-through dynamics from market to retail rates on a panel of euro zone banks for the 1994-2002 period. In the spirit of related literature, an incomplete short-run pass-through is documented. Interest rates on consumer loans and current account deposits tend to be the most rigid. The dynamics of interest rate cycle is asymmetric, with the loan (deposit) rates exhibiting downward (upward) rigidity. The authors argue that strong competition suppresses this anomaly. In a more recent study, Van Leuvensteijn *et al.* (2008) explicitly test for the importance of bank competition in monetary transmission mechanism on a sample of euro zone data. Tighter competition depresses the lending rates, which corroborates findings from the other studies. Somewhat counterintuitive, deposit rates also tend to fall. This result is justified by the fact that, in an attempt to compensate for the reduction in loan market income, banks decrease their deposit rates as well.

The size of mortgage market may approximate the economy's credit dependence. In the context of our analysis, a number of studies have explored its role in monetary transmission mechanism. Using a sample of OECD economies, Sá *et al.* (2011) find that more developed mortgage market amplifies the effects of monetary policy and capital inflow shocks. This is indicative about

[2] The sample comprises Germany, Italy, France, Spain, Netherlands and Belgium.

the presence of financial accelerator. Specifically, in a well-developed market, households can pledge higher value of their residential property as a collateral. Since small interest rate shifts can have a large impact on households' ability to service the debt, risk premium is more sensitive in the case of highly indebted agents. Hence, this type of market setting fosters monetary transmission.

Goodhart & Hofmann (2008) document an operative multifaceted link between monetary policy, credit, housing prices and real economy in OECD economies. More precisely, money growth affects credit and house prices, whereas house prices exert influence over credit and money. The same can be stated about the impact of credit on the remaining two variables. Booming housing market strengthens the repercussions of money and credit shocks on house prices. However, this result must be taken with a pinch of salt due to wide confidence intervals. Assenmacher-Wesche & Gerlach (2008) construct a range of proxies for financial structure, including the importance of floating rate lending, the possibility of mortgage equity withdrawal, the loan-to-value ratio for new mortgages, the mortgage-debt-to-GDP ratio and securitization, to investigate its interaction with the monetary transmission. Impulse responses of residential property prices to a monetary policy shock do hinge upon financial structure, albeit its impact is not overarching. The authors hint that financial structure variables may offset each other, diminishing the difference between the economies.

The final strand of literature deems legal system the crucial factor of monetary transmission, due to its correlation with the financial structure. Importantly, it builds on the findings related to bank lending and balance sheet channel. Cecchetti (1999) attributes cross-country asymmetries in monetary transmission evident in the euro area precisely to divergent legal systems. This study argues that legal systems cause differences in size, concentration and stability of the banking sector, which entails inconsistent responses of inflation and output to monetary policy shocks. Asymmetries are further reinforced by the size and scope of capital markets and availability of alternative sources of financing. Elbourne & De Haan (2004) argue that these findings are not robust across different specifications. Namely, combining different indicators of financial structure into one may produce spurious correlations. When individual indicators are inspected, the correlation with monetary transmission turns out insignificant.

2.3 Evidence on the Credit Channel

Empirical studies offer a certain degree of support for the hypothesis of non-linear propagation of monetary policy shocks. Macro-level studies generally establish the evidence of asymmetries without pointing to a specific transmission channel. On the other hand, studies employing the micro-level data tend to explore either the broad credit channel. We present the empirical evidence and discuss some methodological considerations below.

2.3.1 Macro-level Evidence

Macro-level analyses have been conducted for a wide range of countries. Early studies rely on the stylized facts and simple regressions to detect the asymmetries in monetary transmission. In one of the pioneering studies, McCallum (1991) investigates the effectiveness of monetary policy over the phases of credit cycle. Based on several criteria, including an unusual increase in federal funds rate, high borrowing by the commercial banks, rate inversion, deterioration of business balance sheets and a high demand for short-term credit, the author finds evidence of seven post-1950s credit crunch periods in US. Methodology is based on a linear regression, which includes dummy variable for the state of credit regime (one if economy is credit restricted, zero otherwise). The key findings indicate high potency and substantial role of monetary shocks in GDP fluctuations when credit rationing is active. Galbraith (1996) tests for possible threshold effects of money changes on output in US and Canada. The model takes the following form:

$$y = a_0 + \sum_{i=1}^{p} \alpha_i y_{t-i} + \sum_{i=1}^{s} \beta_i g_{t-i} + \sum_{i=1}^{t} \gamma_i m_{t-i}, \tag{2.2}$$

where y_t is the real GDP, g_t real government expenditure on goods and services, and m_t real M1 aggregate. All the variables are in their logarithmic form. Galbraith employs Monte Carlo simulation to select the threshold value of money supply. The study detects substantial asymmetries. These findings are plagued by a high threshold, which does not allow for clear distinction between the periods of tight and loose monetary policy. Hence, asymmetries may be attributed to some phenomenon other than the credit rationing.

The bulk of macro-level analyses uses some variant of Vector Autoregression (VAR) model, which is designed to capture interdependencies among multiple series. Albeit atheoretical, this model has been fairly successful in replicating some of the key facts of business cycle fluctuations in developed countries. Two influential works, by Christiano *et al.* (1999) and Mojon & Peersman (2001), summarize stylized facts about monetary transmission for US economy and individual countries of the euro area. However, this vein of literature fails to consider the existence of asymmetries inherent to the underlying theory and implicit in identifying assumptions of the model.

Peersman & Smets (2001) remedy for the flaws above by combining VAR with a version of Markov switching model, called *Hamilton filter*. Hamilton filter assumes the existence of two regimes, with the likely switching dates determined endogenously. The analysis is undertaken to investigate monetary policy asymmetries in Germany, Netherlands, Spain, Italy, Austria, Belgium and France for the 1978-1998 period. The model contains a vector of exogenous (world commodity price index, the US short-term interest rate and US real GDP) and endogenous variables (real GDP, consumer prices, nominal three-month short-term interest rate and a real effective exchange rate). Three remarkable findings can be carved out. First, monetary tightening has substantial negative effect on output and prices and vice versa. Second, monetary policy shocks are considerably more potent when the economy is in recession. Finally, the evidence of regime switches is weak: interest rate hikes reduce the probability of staying in a boom only slightly, while no direct switch from boom to recession occurs.

Balke (2001) employs Threshold Vector Autoregression (TVAR) on US data (1960:Q1-1997:Q3) to examine the role of credit in monetary transmission. TVAR is designed to capture regime shifts and asymmetries implied by the theoretical models. Contrary to the Markov switching model, TVAR does not assume random regime shifts. Instead, it splits the sample based on the threshold value. The model is specified as follows:

$$Y_t = A^1 Y_t + B^1(L)Y_{t-1} + (A^2 Y_t + B^2 Y_{t-1})I(c_{t-d} \geq \gamma) + U_t, \qquad (2.3)$$

where $Y_t = [y_t, \pi_t, loan - p, R_t]$ (real GDP, inflation rate, loans to the private sector deflated by GDP deflator and the measure of average nominal interest rate respectively), Y_{t-1} is the lag polynomial matrix, c_{t-d} represents a measure of credit conditions and γ stands for the threshold value. γ is a an indicator variable which takes the value one when $c_{t-d} \geq \gamma$ (high credit growth regime) and zero otherwise. Threshold value between tight and loose credit regime is estimated as in Galbraith (1996). The results reiterate the notion of asymmetric interest rate pass-through at different stages of credit cycle. Furthermore, contractionary shocks appear to be more potent than the expansionary ones. However, if endogenous regime shifts emanate from shocks to variables other than the interest rate (consumption, investment), the evidence of asymmetries largely disappears. Atanasova (2003) replicates Balke's study on a sample of UK monthly data, spanning the 1984-2002 period. This paper employs corporate bond spread as a single measure of credit conditions.[3] An estimated model suggests that the effects of monetary shocks on output growth, inflation and corporate bond spread depend on the state of credit market. No evidence of asymmetric responses to expansionary and contractionary shock is found. This finding supports Balke's conclusion that monetary policy displays symmetric effects once the nonlinearities are controlled for. The study by Calza & Sousa (2005) builds on previous works, applying the TVAR with nonlinear impulse responses on a sample of quarterly euro area data for the 1981-2002 period. This paper finds a clear threshold effect related to the credit conditions. The impact of credit shocks on output and inflation differs over the phases of lending cycle. The most profound response is displayed in a state of constrained lending. When one allows for endogenous regime switches, asymmetries become limited in size. This finding may emanate from nonlinear impulse response function: credit shocks affect other variables in the system, which induces regime shifts. Threshold effects appear to be lower than in US.

The studies by Shen & Chiang (1999), Weise (1999), Lo & Piger (2003), Garcia & Schaller (2002), Mandler (2010) and Galvao & Marcellino (2011) establish two stylized facts concerning the response of output to monetary policy action in US. First, policy actions taken during recessions appear to be more effective than those taken during expansions. Second, the evidence of asymmetry related to the direction of shock is relatively sparse. The former is validated for Australia (Jääskelä 2007) and Canada (Fuchun 2010). Shen & Chiang (1999) hypothesize that the source of these regularities is inflation. More specifically, while monetary transmission under the low inflation regime is in line with the theoretical predictions, high inflation renders it dysfunctional. Mandler (2010) adopts the same line of reasoning, claiming that low inflation is the the key factor of US macroeconomic stability since 1990s. On the other hand, Galvao

[3] The author defines corporate bond spread as the redemption yield on ten-year investment-grade corporate bonds minus equivalent maturity yield on risk-free government debt.

& Marcellino (2011) argue that positive deviation of interest rate from the Taylor rule enhances the propagation of monetary shocks. Weise (1999), Garcia & Schaller (2002) and Lo & Piger (2003) do not relate their analyses to any particular theory, opting to use the output growth as a proxy for economy's position in the business cycle instead.

It is important to note that certain studies do not find support for the stylized facts outlined above. On a sample of post-war US data, De Long & Summers (1988) detect the asymmetric interest rate pass-through based on the direction of shock. This finding is supported by Cover (1988) and Thoma (1994). A peculiarity of all the three studies is the use of money supply as a measure of monetary policy. Weise (1999) argues that the discrepancy in results may emanate from methodological shortcomings in the latter group. For instance, Ravn & Sola (1999) find that Cover's evidence of asymmetry largely disappears when one controls for policy shift at the end of Volcker's years. Thoma's model specification suffers from "impossible restrictions": positive and negative realizations of growth rates of money supply and output enter the model separately. Moreover, the coefficients on output and interest rate variables do not vary with the sign of money supply shock or the state of the economy. As Weise highlights, it would be difficult to find a structural model that supports this range of restrictions.

Monetary transmission asymmetries in developing economies remains a relatively unexplored venue. Due to a strong volatility and structural changes, the exact channels of transmission are difficult to identify. Bordon & Weber (2010) and Çatik & Martin (2012) address the impact of policy shifts on monetary transmission mechanism for Armenia and Turkey, respectively. In the early 2000s, both countries enacted a package of reforms based on the adoption of credible monetary policy framework. The empirical analysis rests on TVAR, which clearly detects two distinct regimes: pre-reform and post-reform. Further, in the case of Armenia, Markov switching model identifies a structural break around the time of policy shift. In both countries, reforms have enhanced the propagation of policy rate, exchange rate and credit shocks, albeit to different extents. Upon the reform, monetary transmission in Turkey came to resemble that of other developed economies. In Armenia, it remains in infancy: policy rate and money supply appear to affect the price level only slightly.

Bathaluddin *et al.* (2012) and Saxegaard (2006) explore the ramifications of excess liquidity on monetary transmission. Their studies focus on Indonesia and Subsaharan Africa respectively.[4] Saxegaard (2006) highlights the necessity to distinguish between precautionary and involuntary liquidity. The former refers to liquid funds held by financial institutions in order to buffer against events with adverse implications, while the latter stands for aggregate excess liquidity.[5] Liquidity held for precautionary purposes is found not to bear an inflationary potential, which renders the sterilization unnecessary. On the other hand, involuntary liquidity carries the risk of excessive lending in the environment of high aggregate demand. Hence, it may reduce the effectiveness of monetary policy in controlling inflation. Bathaluddin *et al.* (2012) apply the TVAR approach on Indonesian sample (2001:M8-2010:M9), detecting the shift from low to high involuntary excess liquidity regime in 2005. The results validate Saxegaard's argument that the leverage of monetary authorities in a high liquidity regime decays.

2.3.2 Micro-level Evidence

Micro-level studies in this domain are often motivated by the *small shocks, big swings* puzzle: the observation that small aggregate shocks cause large fluctuations. Behavior of households, corporate and banking sector may be instructive about the monetary transmission asymmetries.

Evidence on the Bank Lending Channel

Empirical studies on US and euro zone have provided a relatively mixed evidence on the issue. Kashyap & Stein (1994) test for the effectiveness of bank lending channel on a sample of US banks. The channel is found to be operative. Namely, under the conditions of monetary tightening, banks are unable to fully substitute for the deposit drain by external sources of finance. Hence, the loan supply is being restricted. In a more elaborate study, Kashyap & Stein (2000) examine the bank lending channel on a quarterly sample of US commercial banks spanning between 1976 and 2003. This study reaffirms the existence of bank lending channel, finding that the monetary contraction induces deep lending cuts at small and less liquid banks. Methodologically, both studies rely on simple regressions. Kishan & Opiela (2000) investigate federally insured commercial banks for the 1980:Q1-1995:Q4 period, documenting a

[4] The study by Saxegaard (2006) comprises the Central African Economic and Monetary Community (CEMAC) countries, Nigeria and Uganda.

[5] Saxegaard defines excess liquidity as bank reserves deposited in the central bank plus cash for daily operational uses minus minimum reserve requirement.

well-functioning bank lending channel. Concerning the strength and distributional effects of monetary policy shocks, this paper comes up with several novel conclusions. Size and capitalization are found to be important determinants of the response to a monetary shock. Loan supply of small undercapitalized banks stumbles in response to a contractionary shock, while their long-term deposits are completely unresponsive. This bolsters the hypothesis that small undercapitalized banks are unable to raise external funds in the periods of tight monetary policy.

Concerning the euro area, Farinha & Robalo Marques (2001) investigate the functioning of bank lending channel on Portuguese micro-level data. The analysis is conducted on a quarterly sample for the 1990:Q1-1998:Q4 period using the panel cointegration technique. This study pinpoints capitalization as a determinant of bank lending channel. In a large study on euro area, Ehrmann *et al.* (2001) draw a comparison on the role of banks and financial markets in monetary transmission in different economies. We highlight the most relevant findings. First, monetary tightening entails loan supply cuts on both euro area and country level. Liquidity is an essential factor of banks' responsiveness to monetary contraction. Neither size nor capitalization appear to be crucial. This is at odds with US evidence, where the capitalization plays an important role. The difference is ascribed to a lower information asymmetry in the euro area, which is due to the prominent role of governments and banking networks. Huang (2003) finds an overall support for the bank lending channel in UK: monetary tightening compromises banks' ability to provide funding, which affects small, bank-dependent firms more severely. This does not hold for non-bank dependent firms. The evidence above indicates that the banking sector does not behave in a complete accordance with the theoretical predictions.

Evidence on the Balance Sheet Channel

The core analyses on the functioning of balance sheet channel were performed on the sample of US manufacturing firms data, as provided by the Quarterly Financial Report of Manufacturing Firms (QFR). Gertler & Gilchrist (1994) employ a large panel of US manufacturing firms for the period spanning between 1977 and 1991 to examine their behavior under different credit market conditions. Intuitively, agents facing higher agency cost of borrowing should bear the brunt of lending cuts at the onset of recession. As an implication, they are expected to account for a greater decline in economic activity. The results validate this line of reasoning: in the wake of contractionary monetary shock, smaller firms undergo a disproportionate decline in sales and inventories relatively to the large ones.

Bernanke *et al.* (1996) employ the same dataset as Gertler & Gilchrist (1994), concluding that small and bank-dependent firms display a strong procyclicality of sales, inventories and net debt, with the effects of industry adjustment being negligible. An influential study by Kashyap *et al.* (1993) explores the behavior of US financial sector, finding that tighter monetary policy induces lending cuts and a growing issuance of commercial paper. The real effects of monetary tightening are nonlinear: investment changes are concentrated mainly among loan-dependent firms, which do not have an access to the public market. A general conclusion is that the negative effects of credit shocks are borne by small and bank-dependent firms.

Using the sample of US data for the 1964:M1-2001:M12 period, Moody & Taylor (2003) asses the forecasting ability of several financial variables. The nominal term spread had a satisfactory forecasting performance during the 1970s and 1980s only. This relationship broke down in 1960s and 1990s. On the other hand, the spread between investment grade corporate debt and government bond yield, called *high yield spread*, had a substantial predictive power during the 1990s. The authors interpret this as a result of financial market evolution, which changes the information content of different variables. In addition, abnormally high yield has a significant short-term forecasting ability, which implies an economic decay in the aftermath of lending cuts.

The approach above omits the fact that credit conditions are not the sole determinant of corporate spreads. Spreads may exhibit a satisfactory forecasting performance due to the fact that they reflect future expected default probabilities. In order to account for these relationships, Gilchrist *et al.* (2009) employ the Factor-Augmented Vector Autoregression (FAVAR) model. The analysis is conducted on a sample of US non-financial firms for the 1990:M1-2008:M9 period. The panel consists of equity prices and individual firms' credit spreads. FAVAR impulse responses show that the unexpected increases in bond spreads cause large and persistent deviations in output. Moreover, shocks emanating from corporate bond market are found to account for more than 30% of forecast error variance in economic activity at two- to four-years horizon. Overall, these results imply that the credit market shocks are an important factor of economic fluctuations.

2.4 Summary of Findings

From the evidence presented above, several key findings can be carved out. First, interest rate pass-through largely hinges upon credit dependence, competitive pressures from financial sector and availability of alternative sources of financing. Second, the response of output and prices to monetary policy shocks is asymmetric, which is due to three reasons. Interest rate hikes compromise the ability of the banking sector to provide funding. Smaller, less liquid and less capitalized banks tend to be affected more severely, due to the inherent difficulties in raising external funding. Furthermore, in the light of negative monetary impulse, agents are likely to cut their spending more if the credit is already constrained than if it is relatively abundant. Monetary contraction reduces an investment demand; debt obligations of the borrowers increase, depleting their liquidity and depressing the investment. Good times raise the profits and boost the expectations. In this instance, a negative shock is unlikely to induce considerable lending cuts. Henceforth, monetary policy tends to be more potent in turmoil. Further empirical evidence demonstrates that firms respond differently to credit shocks. Thus, the prevailing credit conditions serve as an important nonlinear propagator of shocks.

Chapter 3

Monetary Transmission Asymmetries in the Czech Republic

This chapter attempts to model asymmetries in the Czech monetary transmission. In doing so, we do not bind ourselves to any particular transmission channel. Instead, we adopt a macro-level approach, merely investigating the reaction of aggregate variables to a monetary policy shock under different conditions.

3.1 Literature Survey

Studies investigating monetary transmission in the Czech economy have emerged over the recent years. Albeit few unanimous findings can be documented, we find it instructive to summarize the most relevant ones.

3.1.1 General Evidence

Arnoštová & Hurník (2005) estimate Linear VAR on a sample comprising the 1994-2004 period. A temporary rise in nominal short-term interest rate produces a fall in output with the peak response observed after 5 to 7 quarters. The evidence of price puzzle is found. This anomaly disappears on a subsample excluding the period prior to the inflation targeting. Elbourne & De Haan (2004) restrict their analysis to the same subsample and compare the performance of different estimation strategies. Recursive estimation yields the price puzzle, while non-recursive approach produces responses broadly in line with the economic theory. Borys *et al.* (2009) employ a battery of models on a subsample spanning from 1998, reporting a functional transmission mechanism.

Contractionary monetary policy shock has a negative effect on output and price level, reaching the bottom after roughly one year. No evidence of price puzzle is reported. The authors underline a better performance of models using the output gap instead of output growth. Finally, Havránek *et al.* (2010) investigate the Czech monetary transmission by examining the interaction of financial and macroeconomic variables within the framework of Structural VAR (SVAR). The sample has monthly frequency and encompasses the period from the creation of the euro area till September 2009. A couple of findings are worth highlighting. As regards the effects of monetary action, output responds to monetary tightening in a hump-shaped manner; price puzzle emerges and exchange rate initially overshoots, with the appreciation lasting somewhat less than 20 months. The effects of an increase in credit are intuitive: it boosts the aggregate economic activity and prices, with the maximum response after 7 months. Monetary tightening depresses credit almost instantaneously. Finally, it is found that financial variables tend to improve the forecast of GDP and prices. However, their forecasting performance varies, with the improvement evident over the course of the 2008-2009 crisis.

3.1.2 Specific Aspects

Studies analyzing only one particular aspect of monetary transmission are also insightful. Babetskaia-Kukharchuk (2007) uses a range of models to investigate the exchange rate pass-through in the Czech economy. The author operates with monthly data encompassing the 1996-2006 period. The propagation is relatively fast, with an entire shock being transmitted in the first six months. In absolute terms, the pass-through is rather weak: total response to a shock is around 25% of its magnitude over an entire sample. Horváth (2008b) addresses the issue of asymmetric conduct of monetary policy. Motivation for this type of analysis emanates from the question of credibility. The success of inflation targeting regime in the Czech Republic hinged upon anchored inflation expectations. Hence, target overshooting was a more dangerous outcome than the undershooting. As a result, monetary authorities may have countered the overshooting with a more aggressive policy stance. Estimate of the monetary policy rule over the 1998-2007 sample suggests that, upon the adoption of inflation targeting, Czech National Bank indeed responded more aggressively if inflation forecast was exceeding the target. Evidence of asymmetries vanishes when the same estimation is pursued over the 2002-2007 subsample. Vašíček (2011) augments this analysis with the examination of Minutes of Bank Board meetings. Policy record reveals a greater distaste for inflation. Further, the threshold analysis of monthly data spanning the 1998-2010 period is undertaken. The findings indicate asymmetries in monetary policy stance related to the degree of financial distress.[1] Namely, financial instability appears to prompt a more decisive response from monetary authorities. Contrary to Horváth (2008b), no evidence of asymmetric reactions to the inflation target overshooting/undershooting is reported. Monetary authorities do not respond asymmetrically to downturns and expansions, either.

3.1.3 Evolution

The last vein of literature is concerned with the evolution of monetary transmission mechanism. Darvas (2009) estimates the Czech monetary transmission within the framework of Time-Varying Parameter Vector Autoregression (TVP VAR) model with constant volatility. This approach recognizes gradual changes in the economy through the concept of time-varying coefficients. The results imply somewhat more complete interest rate pass-through over time. Albeit the response of output in 2008 was stronger than in 1996, it must be underlined that the evolution of monetary transmission was not monotonous. Franta *et al.* (2012) employ TVP VAR with stochastic volatility. To account for macrofinancial linkages, the system is augmented with several financial variables. Using the 1996-2010 sample, prices and output are found to be more responsive to monetary policy. Interest rate pass-through was not weakened over the course of 2008-2009 crisis. On the other hand, the impact of exchange rate shocks is found to decline over time, reflecting anchored inflation expectations. Credit shocks had a substantial impact on output and prices during the period of bank restructuring in the early 2000s, remaining negligible during the recent crisis. Weaker impact of credit fluctuations on aggregate economic activity is attributed to greater stability of the Czech financial system.

[1] Financial distress index is constructed as a composite index of five variables: (i) the 12-month rolling beta (derived from the capital asset pricing model - CAPM); (ii) stock market returns (y-o-y change multiplied by minus one, implying that a fall in stock prices leads to an increase in the index); (iii) stock market volatility (six-month rolling monthly squared stock returns); (iv) the sovereign debt spread (the 10-year government bond yield minus the 10-year US Treasury bill yield; (v) the exchange market pressure index (m-o-m percentage change in the exchange rate and total reserves minus gold)

3.2 Methodology

Linear VAR is frequently used in multivariate time series modeling. A plethora of papers on monetary transmission resorts to this particular approach. However, it fails to capture the nonlinear system dynamics, which may produce serious forecasting errors. This chapter puts forward several notions and problems associated with Linear VAR. Building on the discussion on alternative modeling approaches, we proceed to describe the specification used in this study – Logistic Smooth Transition Vector Autoregression (LSTVAR) model.

3.2.1 Linear VAR

Large structural macroeconomic models of 1970s used to be criticized on the grounds of dubious descriptive and forecasting power. Lucas critique brought about their eventual demise. At the time, Sims (1980) pioneered VAR, which later became the most widely applied technique in macroeconomics.

The Basics

VAR is a statistical model used to capture the dynamics and interdependencies among multiple time series. As such, it represents a mere extension of univariate autoregression model.[2] Each of the n variables depends on its own lagged values, current and past values of the remaining $(n-1)$ variables and the white noise. This method performs well in data mimicking and forecasting. Structural inference and policy analysis, however, are tainted by the identification problem, i.e. the necessity to distinguish between correlation and causation (Stock & Watson 2001).

VAR in a bivariate form and with one lag is given by:

$$y_{1t} = b_{10} + b_{12}y_{2t} + \gamma_{11}y_{1,t-1} + \gamma_{12}y_{2,t-1} + \epsilon_{1t} \tag{3.1}$$

$$y_{2t} = b_{20} + b_{21}y_{1t} + \gamma_{21}y_{1,t-1} + \gamma_{22}y_{2,t-1} + \epsilon_{2t} \tag{3.2}$$

This specification is plagued by inconsistency problem, since $Cov(y_1, \epsilon_{2t}) \neq 0$ and $Cov(y_2, \epsilon_{1t}) \neq 0$. Matrix representation of the system above is given by

$$By_t = \Gamma_0 + \Gamma_1 y_{t-1} + \epsilon_t, \tag{3.3}$$

[2] Univariate autoregression is a single-equation, single-variable linear model.

which is labeled *stuctural form VAR*. This specification applies economic theory to recover the original linkages among the variables. Pre-multiplying the equation (3.3) by B^{-1} yields the *reduced form VAR*:

$$y_t = A_0 + A_1 x_{t-1} + e_t, \tag{3.4}$$

where $A_0 = B^{-1}\Gamma_0$ and $A_1 = B^{-1}\Gamma_1$ and the relationship between reduced and structural error terms is given by $e_t = B^{-1}\epsilon_t$. Correlation among the variables produces serial correlation among the shocks in u_t as well.[3] Hence, OLS estimators are no longer Best Linear Unbiased Estimators (BLUE). It should be noted that the reduced form is a mere vehicle to summarize dynamic properties of the data.

Interpretation of the model requires structural form, which tends to be overparameterized. The system (3.1) – (3.2) contains 10 unknowns (8 coefficients and 2 variances of error terms). Reduced form, on the other hand, gives us only 8 parameters (6 coefficients, 2 variances and the covariance of the error terms). Hence, certain identifying restrictions are necessary. In a bivariate specification, matrix B from equation (3.3) will be of 2×2 dimensions; the elements on its diagonal are ones, so B has $2 \times 2 - 2 = 2$ unknown parameters. There are also 2 unknown variances of ϵ, so the total number of unknowns is 4. From the estimation of variance-covariance matrix of e_t, we can find 3 parameters; one remains unknown, which is the number of restrictions to be imposed.[4] Restrictions are typically imposed on the vector of residuals to find the true shocks which are independent.[5] The identifying restriction $b_{12} = 0$ yields the so called *recursive form VAR*. Assuming that y_{1t} stands for inflation, while y_{2t} is the policy rate, recursive form interpretation is straightforward: inflation is affected by historical values only, while the policy makers account for a contemporaneous value of inflation.

[3] For the proof, please refer to Enders (2004).

[4] In a K-variable VAR, the number of necessary restrictions is the difference between the number unknowns, K^2 and the number of parameters estimated from variance-covariance matrix, $\frac{K^2+K}{2}$ which equals to $\frac{K^2-K}{2}$.

[5] This is the *recursiveness assumption*, originally proposed by Sims (1980).

We now indicate the recursiveness assumption in a three-variable specification. Rewrite the equation (3.3) in a matrix form:

$$\begin{bmatrix} 1 & b_{12} & b_{13} \\ b_{21} & 1 & b_{23} \\ b_{31} & b_{32} & 1 \end{bmatrix} \begin{bmatrix} y_{1t} \\ y_{2t} \\ y_{3t} \end{bmatrix} = \begin{bmatrix} b_{10} \\ b_{20} \\ b_{30} \end{bmatrix} + \begin{bmatrix} \gamma_{11} & \gamma_{12} & \gamma_{13} \\ \gamma_{21} & \gamma_{22} & \gamma_{13} \\ \gamma_{31} & \gamma_{32} & \gamma_{33} \end{bmatrix} \begin{bmatrix} y_{1,t-1} \\ b_{2,t-1} \\ b_{3,t-1} \end{bmatrix} + \begin{bmatrix} \epsilon_{1t} \\ \epsilon_{2t} \\ \epsilon_{3t} \end{bmatrix}$$

The relationship between the matrix of coefficients in reduced and structural form is given by $A_1 = B^{-1}\Gamma_1$:

$$\begin{bmatrix} 1 & a_{12} & a_{13} \\ a_{21} & 1 & a_{23} \\ a_{31} & a_{32} & 1 \end{bmatrix} = \begin{bmatrix} 1 & b_{12} & b_{13} \\ b_{21} & 1 & b_{23} \\ b_{31} & b_{32} & 1 \end{bmatrix}^{-1} \begin{bmatrix} \gamma_{11} & \gamma_{12} & \gamma_{13} \\ \gamma_{21} & \gamma_{22} & \gamma_{13} \\ \gamma_{31} & \gamma_{32} & \gamma_{33} \end{bmatrix}$$

Assuming that the equation (3.3) in trivariate form captures monetary policy decision-making, it will contain inflation, policy rate and unemployment, ordered in this particular sequence. This ordering scheme implies that the contemporaneous values of inflation enter the policy rate equation, whereas unemployment appears with a lag only. Identification restrictions $a_{12} = a_{13} = a_{23} = 0$ yield the recursive form VAR. Matrix A takes the form below:

$$A = \begin{bmatrix} 1 & 0 & 0 \\ a_{21} & 1 & 0 \\ a_{31} & a_{32} & 1 \end{bmatrix}$$

Two zeros in the first row indicate that the monetary policy shocks are independent from shocks to variables contained in the inflation equation. To rephrase, monetary policy shock affects inflation with a lag only. A zero in the the second row implies that monetary authorities do not observe the actual unemployment rate at the time of decision-making. Decomposing structural terms in a triangular fashion is called *Cholesky decomposition.*

Stability

An important issue in VAR model is stability. Consider the following process:

$$y_t = A_0 + A_1 y_{t-1} + e_t, \tag{3.5}$$

which after n backward iterations reduces to:

$$y_t = (I + A_1 + A_1^2 + ... + A_1^n)A_0 + A_1^{n+1}x_{t-n-1} + \sum_{i=0}^{n} A_1^i e_{t-i}. \tag{3.6}$$

Depending on the value of A_1, three possible scenarios emerge:

- Shocks gradually die away: $|A_1| < 1 \implies |A_1^n| \to 0$ as $n \to \infty$,
- Shocks never die away: $|A_1| = 1 \implies |A_1^n| = 1\ \forall N$, in which case the Equation (3.5) transforms into $y_t = y_0 + \sum_{i=0}^{\infty} u_t$ as $N \to \infty$,
- Shocks magnify over time: $|A_1| > 1 \implies |A_1^n| > |A_1^{n-1}| > ... > |A_1|$.

The model is stable only in the first case, i.e. all eigenvalues of A_1 lie inside the unit circle. Thus, VAR requires stationary time series. However, Sims (1980) cites the danger of information loss as the case against differencing. More specifically, he argues that VAR is created to capture the time series dynamics and interaction between the variables, not to estimate the parameters. Further, differencing obscures the co-movements among the series, i.e. omits possible cointegrating relationships. On the other side of token, non-stationary series may induce false relationships and spurious correlations. We decide against the dogmatic approach, treating this issue on a case-to-case basis.

One of the key tenets of VAR analysis is invertibility condition (Wold representation). Consider the equation:

$$y_t = \sum_{i=1}^{\infty} \psi_i \epsilon_{t-i}, \tag{3.7}$$

which ensures that the forecast for any stationary process is produced by the means of weighted average of past forecast errors. In lag notation, the equation (3.7) reads

$$y_t = (1 + \psi L)^{-1} \epsilon_t \tag{3.8}$$

Under invertibility condition, (3.8) can be rewritten as

$$\epsilon_t = y_t - \psi y_{t-1} + \psi^2 y_{t-2} + ... \tag{3.9}$$

The equation (3.9) holds if $|\psi \leq 1|$, implying that the innovation term ϵ_t is a function of the past values of observables y_t. In the case of non-invertibility, the innovations from VAR will not match the shocks and pertinent responses from an economic model.

Linear Impulse Response Function

Interpretation of structural VAR is hindered by a large number of coefficients. Impulse response function (IRF) is a simple way of tracking the time path of various shocks on the variables contained in VAR system (Enders 2004). To illuminate the concept, consider the recursive form VAR. One off-unit shock to ϵ_{2t} does not have a contemporaneous effect on y_t, but is being transmitted into ϵ_{1t+1}. Hence, contemporaneous shock in policy rate affects the inflation in $(t+1)$. Impulse response depicts the scale of this impact over a certain horizon. In mathematical notation, IRF can be represented as the difference between the two realizations of a conceptual experiment. In the first one, the system is hit by a random shock at time t; whereas in the second, no shocks occur in the time interval between t and $(t+n)$ (Koop *et al.* 1996):

$$I_y(n, \epsilon, \omega_{t-1}) = E[y_{t+n}|V_t = \epsilon, ..., V_{t+n} = 0, \omega_{t-1}] - E[y_{t+n}|V_t = 0, ..., V_{t+n} = 0], \tag{3.10}$$

where V_t is a random shock, ω_{t-1} represents the history of y_{t-1} up to $(t-1)$ and $n = 1, 2, 3, ...$

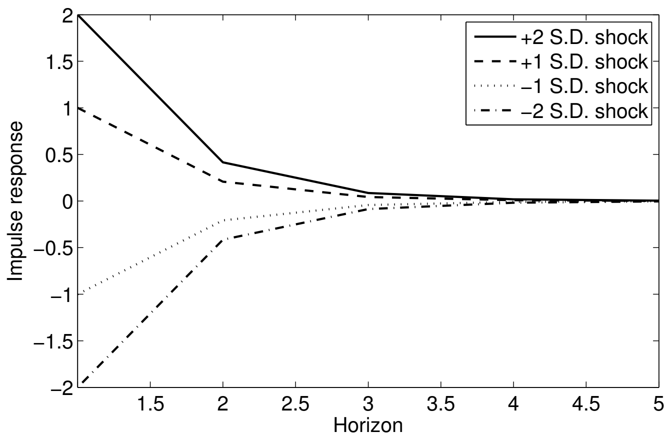
Linear IRFs are symmetric and history independent, $\omega_{t-1} = 0$. Wold decomposition represents the innovation term as a combination of lagged observables; hence the assumption of zero shocks in intermediate periods. Different values of ϵ only rescale this measure, implying that the responses do not depend on the sign or the size of shocks nor their history. Figure 3.1 illustrates the impulse responses of AR(1) process to different values of an initial shock.[6] It can be inferred that the responses are symmetric in all instances.

IRFs are ill-suited to study the systems with changing dynamics. Nonlinear models do not have a Wold representation; hence, the assumption of zero shocks in intermediate periods would lead to inaccurate inferences. Further, the reaction of regime switching systems to the same shock hinges upon the initial conditions. As an example, consider the Threshold Autoregression (TAR):

$$\Delta y_t = A_1 \Delta y_{t-1} + A_2 \Delta y_{t-1} + I(\Delta y_{t-1} \geq 0) + V_t, \tag{3.11}$$

[6] The AR(1) process $y_t = 0.2 y_{t-1} + \epsilon_t$ was generated in EViews.

Figure 3.1: Impulse responses, AR(1)



Source: author's computations.

where $I(\Delta y_{t-1} \geq 0) = 1$ if $\Delta y_{t-1} \geq 0$ and zero otherwise. Define $A = A_1 + A_2$, where $0 \leq A \leq 1$ and $0 \leq A_1 \leq 1$. For a particular history Δy_{t-1} and $V_t = \epsilon$, two sets of impulse responses emerge:

(a) $I_y(n, \epsilon, \Delta y_{t-1} = 0) = \epsilon \frac{1-A^{n+1}}{1-A}$ if $\epsilon \geq 0$

(b) $I_y(n, \epsilon, \Delta y_{t-1} = 0) = \epsilon \frac{1-A^{n+1}}{1-A_1}$ if $\epsilon < 0$

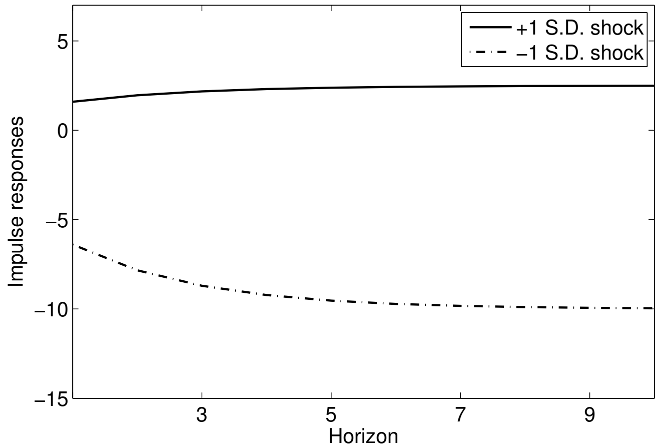
Figure 3.2 shows the impulse responses of TAR(1) with the coefficients $A_1 = 0.9$ and $A_2 = 0.3$. The responses to positive and negative one-standard deviation shocks are clearly asymmetric and of different magnitudes.

Linear impulse response fails to describe the effects of mutually correlated shocks, which is referred as the *composition effect property*. In the system (3.1) – (3.2), ϵ_{1t} and ϵ_{2t} are correlated, which induces contemporaneous effect of ϵ_{1t} on both y_{1t} and y_{2t}. Hence, the responses depend on the composition of shocks, rendering linear IRF inadequate.

Criticism

The appeal of VAR as an empirical tool vanishes in theory. Critics have articulated its inability to recover the true structural relationships as the most serious flaw. Canova (2006) argues that conventional identification restrictions are created merely to facilitate an econometric calculation. As such, they are disassociated from the structural models used for interpretation. On the other

Figure 3.2: Impulse responses, TAR(1)



Source: author's computations.

hand, capturing a variety of structural relationships accurately is daunting. Proper VAR specification should reflect the theory as closely as possible, without catering to the econometric convenience.

Fernández-Villaverde *et al.* (2005) argue that potential non-invertibility may discredit the model by generating large prediction errors. This case usually applies if some variables are not observed. Briefly, let the equilibrium of an econometric model have the following state space representation:

$$x_{t+1} = Ax_t + Bw_{t+1} \tag{3.12}$$

$$y_{t+1} = Cx_t + Dw_{t+1} \tag{3.13}$$

where x_t is the vector of states, y_t is the vector of observable variables, w_t is the vector of economic shocks and D is a square matrix of full rank. The model is invertible if the absolute values of eigenvalues of $A - BD^{-1}C$ are strictly less than one.[7] If invertibility condition fails, the impulse responses from state space model do not match those from VAR representation.

SVAR models in finite samples may be unable to recover the true system dynamics, even if the correct identifying restrictions are employed. This argument has been put forward by Chari *et al.* (2008), who find that SVAR fails to recover impulse responses to technology shocks on the data generated from multiple-shock business cycle models. Insufficient number of lags is cited as the core obstruction to the model's credibility. Including higher number of lags, on

[7] For the proof, please refer to Fernández-Villaverde *et al.* (2005)

the other hand, entails practical shortcomings: the model becomes highly parameterized, which requires long time series. Moreover, the presence structural changes renders the parameter estimates unstable.

3.2.2 Selected Nonlinear Models

Unsatisfactory performance of Linear VAR in a variety of applications led to the development of new econometric techniques. Before describing the Logistic Smooth Transition Vector Autoregression (LSTVAR) model, we present the alternative nonlinear techniques which have been used to investigate the monetary transmission mechanism.

Markov Switching Model

One of the initial attempts to address the issue of regime changes was Markov switching model. System dynamics oscillates between the s_1, s_2,..., s_n states. State switches are governed by the *Markov process*:

$$P(y_{t+1} = s_i | y_1, y_2, ..., y_t) = P(y_{t+1} = s_i | y_t), \quad i = 1, ..., n \tag{3.14}$$

The equation (3.14) implies that the conditional probability distribution of future states hinges only upon the current state, i.e. is conditionally independent upon the past values. Hence, the forecast is based upon the current state and the transition probability matrix:

$$A = \begin{bmatrix} P_{11} & P_{12} & ... & P_{1m} \\ P_{21} & P_{22} & ... & P_{2m} \\ \vdots & \vdots & \ddots & \vdots \\ P_{m1} & P_{m2} & ... & P_{mm} \end{bmatrix}$$

where $A_{ij} = P(y_{t+1} = s_j | y_t = s_i)$ is the transition probability between the regimes i and j and $\sum_{j=0}^{m} P_{ij} = 1$. The switches occur due to the movements of unobservable variable. Parameters are estimated using the maximum likelihood procedure. As far as our study is concerned, the model assumptions are at odds with the dynamics of the Czech economy. More specifically, the structural changes can be related to a particular sequence of events, which contrasts to random state switches.

Threshold Vector Autoregression (TVAR)

In TVAR, regime switches are governed by the motion of selected state variable. A variation of this model, LSTVAR, allows for smooth transition between the regimes. Dataset is split based on the threshold, with the impulse responses computed for each subsample. This estimation strategy may have problematic implications. Consider the model of economy with output, prices, interest rate and credit. We may want to inspect the effectiveness of monetary policy in expansion and recession separately. Accordingly, the impulse responses are computed both for the recession and expansion subsample. However, TVAR dismisses the time dimension, as the reaction of the economy to monetary shocks may change over time. Thus, the estimates may be imprecise.

Time-Varying Parameter Vector Autoregression (TVP VAR)

Time-varying parameter Vector Autoregression (TVP VAR) model is a relatively novel approach to monetary transmission modeling, emanating from the development of Bayesian econometrics.[8] Parameters of interest are taken as the true values of random variables. Consequently, Bayesian econometrics focuses on their distributional properties. As of the TVP VAR, parameters are allowed to change with each observation, usually following the random walk assumption. This estimation strategy may render the model seriously overparameterized.

One of the initial attempts to model the monetary transmission within the TVP VAR framework concerned the bad luck vs. bad policy question (Cogley & Sargent 2001). The core shortcoming of this study is the assumption of constant volatility of shocks, which neglects possible heteroskedasticity of shocks and nonlinear relationships among the model variables. Later studies, most notably by Cogley & Sargent (2005) and Primiceri (2005), allow for stochastic volatility of shocks and system nonlinearities. However, TVP VAR demonstrates the evolution of impulse responses over time, without explicitly indicating the impact of certain variables on monetary transmission.

3.2.3 The Basics of LSTVAR

Our model specification pertains to Logistic Smooth Transition VAR (LSTVAR). The dynamics of LSTVAR oscillates between two different states. Tran-

[8] Readers interested in this topic should consult Koop (2003).

sition between the regimes occurs at a certain speed rather than abruptly. Impulse responses are computed for each regime. This section presents our model specification and outlines the pertinent simulation procedure.

The Model

In line with Weise (1999), we set up the following model:

$$X_t = X_0 + A(L)X_{t-1} + (\theta_0 + \theta(L)X_{t-1})F(z_t) + u_t, \tag{3.15}$$

where X_t is an $n \times 1$ vector of endogenous variables, $A(L)$ and $\theta(L)$ are $n \times n$ m-th order polynomials in the lag operator, u_t are unobservable shocks with variance-covariance matrix Ω_t for $t = 1, ..., T$; z_t is the switching variable, while $F(z_t)$ represents logistic function which governs the transitions between the regimes. $F(z_t)$ is assumed to take the following form:

$$F(z_t) = (1 + exp[-\gamma(z_t - c)/\sigma_z])^{-1} - 1/2, \gamma > 0, \tag{3.16}$$

where c is the threshold around which the system dynamics evolves, γ is the "smoothness parameter", while σ_z represents the standard deviation of the switching variable. Function $F(z_t)$ is bounded between zero and one. In the limit, four possible combinations emerge:

(a) $(z_t - c) \to -\infty \implies F(z_t) \to 0 \implies$ linear model
(b) $(z_t - c) \to +\infty \implies F(z_t) \to 1 \implies$ linear model
(c) $\gamma \to 0 \implies F(z_t) \to 1 \implies$ linear model
(d) $\gamma \to \pm\infty \implies$ threshold autoregression with two regimes.

The parameter σ_z normalizes deviations of the z_t from the threshold value and is calculated as the standard deviation of the switching variable.

Pre-Testing for Nonlinearity

Prior to the actual estimation, one must establish whether the hypothesis of linearity can be rejected in favor of nonlinear specification. The usual approach is to conduct a grid search for potential threshold values, each time estimating the selected model specification and computing the Lagrange Multiplier (LM) test statistics on the restriction of equality between linear and nonlinear models. Potential threshold values are selected from parameter space $\Gamma = [z, \hat{z}]$, where Γ is restricted to a sample range of the switching variable. Equation (3.15) is then estimated for all the values from Γ.

LM test posits the null hypothesis of coefficients' stability across the sample, $H_0 : \gamma = 0$, against the alternative $H_1 : \gamma > 0$.[9] Testing for nonlinearity in each equation separately is conducted in three steps. Assuming that our initial model is the k-variable VAR with p-lags, where $W_t = (X_{1t-1}, X_{1t-2}, ..., X_{1t-p}, X_{2t-1}, ..., X_{kt-p})$, the nonlinearity test is performed as follows (Weise 1999):

(a) Calculate the restricted sum of residuals from the regression of explanatory variables on their lagged values

$$X_{it} = \beta_{i0} + \sum_{j=1}^{pk} \beta_{ij} W_{jt} + u_{it},$$

and denote $SSR_0 = \sum_{t=1}^{p} \hat{u}_{it}^2$.

(b) Calculate the unrestricted sum of residuals from the regression of residuals on lagged values of explanatory variables

$$u_{it} = \alpha_{i0} + \sum_{j=1}^{pk} \alpha_{ij} W_{jt} + \sum_{j=1}^{pk} \delta_i z_t W_{jt} + v_{it}$$

and denote $SSR_1 = \sum_{t=1}^{p} \hat{v}_{it}^2$.

(c) Compute the test statistics

$$LM = T(SSR_0 \cdot SSR_1)/SSR_0,$$

where T stands for the number of observations. Detecting nonlinearities in a system of equations requires the application of log-likelihood test. Test formulation is the same as previously: $H_0 : \gamma = 0$ versus $H_1 : \gamma > 0$ in all the equations. Denote the estimated variance-covariance matrices of residuals from restricted and unrestricted regressions as $\Omega_0 = \sum \hat{u}_t \hat{u}_t'/T$ and $\Omega_1 = \sum \hat{v}_{it} \hat{v}_{it}'/T$ respectively. The test statistics is defined as $LR = T(log|\Omega_0| - log|\Omega_1|)$.

The testing procedure in both cases is tainted by a nuisance parameter problem, as the threshold parameter is not identified under H_0. Hence, the test statistics conforms to probability distribution different than the χ^2. The bootstrap procedure to approximate the asymptotic distribution in a given scenario is implemented in two steps. First, construct the artificial series by estimating a baseline linear model and taking n draws from the residuals, where

[9] Testing for $\gamma = 0$ entails an identification issue. In this case, γ would also take the negative values. The shape of the transition function would switch around c, while the parameters in $\theta(L)$ would simply adjust.

n is the desired length of series; second, compute the F or LR statistics for each of the series and comparing them to the actual data. The p-value is calculated as the number of times the F or LR statistics from artificial series exceeded the actual one, divided by one thousand (Weise 1999).

Optimal Threshold Value

If the linear specification is rejected, optimal threshold value is selected from the aforementioned grid search. In order to ensure that the adequate number of time series observations lies on each side of the threshold, the highest and the lowest 15% observations are eliminated from the search. Upon estimating the equation (3.15), optimal threshold and smoothness parameter are selected as a combination of values which minimizes the log-determinant of variance-covariance matrix of residuals S:

$$z_{opt} = \arg\min S(z), \tag{3.17}$$

where $z \in \Gamma$.

Generalized Impulse Response Function (GIRF)

Shortcomings of linear IRF led to the development of its generalized version. Generalized impulse response function (GIRF) is designed to capture the non-linear system dynamics. Responses are conditioned on the sign, the size and the history of shocks, while the endogenous regime switches throughout the duration of response are allowed. Formal definition of GIRF from the equation (3.10) applies, with the assumption $\omega_{t-1} \neq 0$.

Given that the nonlinear model is already specified, GIRF can be computed using the Monte Carlo integration technique. The crux of this method is in simulating different histories and shocks R times and taking an average. Koop *et al.* (1996) propose the following procedure:

1. Pick a history ω_{t-1} and a random shock $V_t = \epsilon$. For ω_{t-1}, the actual values of lagged endogenous variables at a particular date are used.

2. For a given horizon N, randomly select $(N + 1) \times R$ values of K-dimensional shock. The shocks are drawn with replacement from estimated residuals of the model and assumed to be jointly distributed within one period. Hence, if the date t shock is drawn, all residuals for that particular date are

collected.

3. Simulate the evolution of endogenous variables: given the initial conditions V_t and ω_{t-1}, use the first N shocks to compute the realization y^0_{t+n} for $n = 1, ..., N$.

4. Use the same draw of N random shocks plus one additional draw to produce a realization y^0_{t+n} of the time series for $n = 1, ..., N$ dependent on the initial condition ω_{t-1}.

5. Repeat steps 3 and 4 R times and form the averages of each individual component:

$$\overline{y}_{R,t+n}(V_t, \omega_{t-1}) = \frac{1}{R}\sum_{i=0}^{R-1} y^i_{t+n}(V_t, \omega_{t-1}), n = 1, ..., n \tag{3.18}$$

$$\overline{y}_{R,t+n}(\omega_{t-1}) = \frac{1}{R}\sum_{i=0}^{R-1} y^i_{t+n}(\omega_{t-1}), n = 0, ..., n \tag{3.19}$$

By the Law of Large Numbers, these averages will converge to the conditional expectations, $E[Y_{t+n}|v_t, \omega_{t-1}]$ and $E[Y_{t+n}|\omega_{t-1}]$, required in the definition of GIRF.

Simulation procedure poses the question of the nature of shocks. In our original specification, we assumed that $V_t = \epsilon$ is a random shock, i.e. the realization of the same stochastic process that generates y_t, while the variables generating history ω_{t-1} are arbitrary. Alternatively, the variables y_{t-p} may be assumed random, when we denote history Ω_{t-1}. GIRF can be constructed as one of the following combinations:

(a) Arbitrary current shock and arbitrary history:

$$GI_y(n, \delta, \omega_{t-1}) = E[y_{t+n}|v_t = \delta, \omega_{t-1}] - E[y_{t+n}, \omega_{t-1}], \tag{3.20}$$

for n=0,1,...

(b) Random current shock and random history:

$$GI_y(n, \delta, \omega_{t-1}) = E[y_{t+n}|V_t = \delta, \Omega_{t-1}] - E[y_{t+n}, \Omega_{t-1}], \tag{3.21}$$

(c) Arbitrary current shock and random history:

$$GI_y(n, \delta, \Omega_{t-1}) = E[y_{t+n}|v_t =, \Omega_{t-1}] - E[y_{t+n}, \Omega_{t-1}]. \tag{3.22}$$

3.3 Empirical Model of the Czech Economy

This section lays out the empirical model of the Czech economy. As part of our estimation strategy, we consider various filtering methods and identification schemes. An extensive description of the dataset is provided, along with the corresponding transformations.

3.3.1 Filtering Method

Prior to the actual estimation, each series has been detrended using the Hodrick-Prescott (HP) filter. HP takes the following form:

$$min \sum_{j=1}^{N} (y_j - g_j)^2 + \lambda \sum_{j=2}^{N-1} (g_{j+1} - 2g_j + g_{j-1})^2, \tag{3.23}$$

where y_t is a cyclical component, defined as the difference between the original signal x_t and the smooth growth component g_t. The first term is a measure comparable to "goodness-of-fit", while the second term measures the "degree of smoothness". Variable λ determines the trade-off between the two terms. The respective values of $\lambda = 1600$ and $\lambda = 14400$ appear to be standard for quarterly and monthly frequency data. Using the local linear trend model with stochastic cycle and calibrated volatilities of components, Harvey & Jaeger (1993) demonstrate that this particular value was tailored to approximate the US business cycle. Specifically, it can be shown that the standard value $\lambda = 1600$ corresponds to 39.4 months, which is a standard duration of US business cycle.[10] Hence, uncritical application of the standard values may lead to inaccurate inferences.

As of the Czech Republic, there is a great deal of uncertainty surrounding the length of a cycle. The studies addressing this issue are scarce. Based on the analysis of stylized facts, Hloušek (2006) finds that, since the beginning of 1990s, at most two cycles can be identified. Poměnková & Maršálek (2011) attempt to detect the cycle via turning points. However, the range of estimates is wide: consumption cycle is found to last between 3.5 and 6.5 years, industrial production between 1.75 and 7.75, while that of GDP is within 3.75 and 11.5 years. Albeit uncritical, we opt for the standard value ($\lambda = 14400$) to avoid bias and ensure comparability with the existing studies.

[10] For the proof, please refer to Kowal (2005).

3.3.2 Identification

Following Mojon & Peersman (2001), we specify our benchmark model as follows:

$$KY_t = A(L)Y_{t-1} + B(L)X_t + \mu_t, \tag{3.24}$$

where Y_t is the vector of endogenous variables and X_{t-1} the vector of foreign exogenous variables. Pre-multiplying the equation (3.24) by K^{-1} yields the reduced form:

$$Y_t = A(L)^* Y_{t-1} + B(L)^* X_t + \epsilon_t, \tag{3.25}$$

where $A^* = K^{-1}A$, $B^* = K^{-1}B$ and $\epsilon = K^{-1}\mu$. ϵ stands for reduced form residuals, which will be used to recover the original shocks μ. The economy is divided into domestic and foreign block, containing p_1 and p_2 variables respectively. Foreign block is included in order to isolate the effects of foreign shocks on economy and therefore, to avoid spurious correlations. In matrix notation, the equation (3.25) reads

$$\begin{bmatrix} Y_{1,t} \\ Y_{2,t} \end{bmatrix} = \begin{bmatrix} A_{11}(L) & A_{12}(L) \\ A_{21}(L) & A_{22}(L) \end{bmatrix} \begin{bmatrix} Y_{1,t-1} \\ Y_{2,t-1} \end{bmatrix} + \begin{bmatrix} B_{11}(L) \\ B_{22}(L) \end{bmatrix} X_t + \sigma \begin{bmatrix} \epsilon_{1,t} \\ \epsilon_{1,t} \end{bmatrix}$$

where $A_{ij}(L)$ takes the dimensions $p_1 \times p_2$. The study by Mojon & Peersman (2001) examines the effects of monetary policy in different euro zone economies. It is assumed that shocks emanating from small open economies (Austria, Belgium and Netherlands) do not affect the EMU-wide fluctuations, while the inverse relationship holds. Accordingly, the variables associated with EMU economic activity become endogenous, but with an imposed block exogeneity restriction. This restriction postulates that domestic endogenous variables do not enter the foreign exogenous block either contemporaneously or with a lag ($A_{21} = 0$). It has been applied in a number of studies on small open economies, most notably by Cushman & Zha (1997) who explore the interactions between US and Canadian economy. Further, it is routinely employed in the papers on monetary transmission in CEE region (Maćkowiak 2005), (Horváth 2008a), (Havránek *et al.* 2010).

In our baseline specification, vector of endogenous variables contains industrial production, net price index, three-month money market rate, exchange rate and total credit to residents and nonresidents:

$$Y_{1,t} = \begin{bmatrix} ip_t & pr_t & ir_t & er_t & cr_t \end{bmatrix}$$

while the exogenous block includes euro area output, aggregate price level and euro area short term interest rate:

$$X_{1,t} = \begin{bmatrix} ei\_gdp_t & ei\_pr_t & ei\_ir_t \end{bmatrix}$$

It needs to be emphasized that the use of GDP *per se* in VAR analysis bears a number of disadvantages. Giordani (2001) argues that the omission of output gap as a forward-looking variable may spuriously produce the price puzzle. Specifically, output gap enters the monetary policy rule, while the data on output is not available at the time of monetary decision-making. Employing the *ex post* series may therefore lead to incorrect identification of the response of monetary policy variable to all the other shocks. Hence, we employ cyclical component of GDP.

The crux of correct identification of monetary shocks is the identification scheme. In a benchmark study on VAR analysis, Sims (1980) proposes recursive ordering and Cholesky identification scheme. Some studies argue that this particular scheme may not be entirely plausible for a small open economy. Assuming that the variables are ordered as in Y_t and X_t, we combine the approaches proposed by Sims & Zha (1998) and Mojon & Peersman (2001):

$$\begin{bmatrix} \epsilon_t^1 \\ \epsilon_t^2 \\ \epsilon_t^3 \\ \epsilon_t^4 \\ \epsilon_t^5 \\ \epsilon_t^6 \\ \epsilon_t^7 \\ \epsilon_t^8 \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ k_{21} & 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ k_{31} & k_{32} & 1 & 0 & 0 & 0 & 0 & 0 \\ k_{41} & k_{42} & k_{43} & 1 & 0 & 0 & 0 & 0 \\ k_{51} & k_{52} & k_{53} & k_{54} & 1 & 0 & 0 & 0 \\ k_{61} & k_{61} & k_{63} & k_{64} & 0 & 1 & k_{67} & 0 \\ k_{71} & k_{72} & k_{73} & k_{74} & k_{75} & k_{76} & 1 & 0 \\ k_{81} & k_{82} & k_{83} & k_{84} & k_{85} & k_{86} & k_{87} & 1 \end{bmatrix} \begin{bmatrix} \mu_t^{ei_{gdp}} \\ \mu_t^{ei_{pr}} \\ \mu_t^{ei_{ir}} \\ \mu_t^{ip} \\ \mu_t^{pr} \\ \mu_t^{ir} \\ \mu_t^{er} \\ \mu_t^{cr} \end{bmatrix}$$

The interpretation of the identification scheme above is essentially the same as in the case of Cholesky decomposition. However, several important changes emerge. First, monetary authorities react to exchange rate fluctuations ($k_{67} \neq 0$). Borys *et al.* (2009) make the same assumption in the study on the Czech monetary transmission. Second, monetary authorities are assumed to respond contemporaneously to changes in industrial production ($k_{64} \neq 0$). A rationale behind is that the excess industrial production indicates excess demand, which prompts monetary authorities to react (Borys *et al.* 2009). Finally, contemporaneous reactions to the price level are ruled out, as the deviation of forecast from the target actually enters the monetary policy reaction function. The first three rows are related to the conduct of euro area monetary policy.

Our benchmark model is similar to that in Havránek *et al.* (2010), with a slightly modified identification scheme. However, our dataset is longer and spans until 2012:M3. The initial exercise, therefore, is to estimate the model over a longer sample, replacing GDP with the industrial production. Next, we evaluate the impact of monetary shocks using generalized impulse responses. This strategy should be more informative, as GIRF is insensitive to ordering and allows for endogenous shocks throughout the duration of the response.

3.3.3 Caveats

The use of this particular identification scheme in LSTVAR has an important caveat: overparameterization. In the two regime specification, the number of estimated parameters doubles in comparison to Linear VAR. Hence, for a model with eight variables and two lags, the number of estimated parameters would be $8 \times 2 \times 2 = 32$. Our dataset contains 171 observations, which would imply a substantial loss of the degrees of freedom. Thus, we need to specify the model in a more parsimonious manner.

Studies employing some version of TVAR generally resort to four- or five-variable specification. Three and four variables are typically used in the papers analyzing the threshold effects in large economies. Weise (1999) employs the first differences of logs of industrial production index, consumer price index and M1 on a US sample spanning the 1960:Q2-1995:Q2 period. This specification is convenient, albeit impractical for our purposes. First, exchange rate omission would be a serious error, due to its relevance in a small open economy. Second, the use of M1 as a measure of monetary policy stance is flawed, as the CNB's main tool is the two-week repo rate. Furthermore, M1 is not indicative about the credit demand, which would taint our estimation procedure. Next, Atanasova (2003) employs a four-variable specification on UK sample (1984:M1-2002:M4), including the annual growth of industrial production index, annual change of the retail price index, M2 and the measure of credit market conditions. The use of M2 as a monetary policy variable is flawed as well. M2 is to a large extent determined by the demand for credit, which induces endogeneity in monetary policy reaction. Thus, monetary policy shifts cannot be detected unambiguously. Atanasova provides justification that the three-month money market rate produces the price puzzle. Contrary to this, benchmark studies on the Czech monetary transmission mechanism find no evidence of the price puzzle. Hence, the use of M2 would be unreasonable. Some studies acknowledge the downsides of this particular strategy, simultaneously employing the interest rate and some measure of monetary aggregate (Karamé & Olmedo 2002; Shen & Chiang 1999). More recent studies by Fuchun (2010) and Jääskelä (2007) on Canada and Australia respectively employ a parsimonious four-variable specification, containing output growth, price level, some measure of credit market conditions and interest rate.

Studies on the threshold effects in small open economies generally employ five variables. Standard model variables comprise output growth, price level and interest rate. Importantly, neither of the studies available omits the exchange rate. The papers by Bordon & Weber (2010) and Bathaluddin *et al.* (2012) on Armenia and Indonesia respectively estimate the regime switching models without exogenous variables. Saxegaard (2006) adopts similar specification for Subsaharan Africa, employing an exogenous variable only if its economic relevance is overarching. For instance, oil price and aid-to-GDP ratio are included as exogenous variables for Nigeria and Uganda respectively.[11] Contrary to these, Çatik & Martin (2012) employ a five-variable baseline specification and three exogenous variables (oil price, US output and US interest rate) when analyzing the monetary transmission in Turkey. However, their estimation sample spans a relatively long period (1986:M1-2010:M11), rendering the inclusion of the aforementioned vector plausible.

3.3.4 LSTVAR Specification

We now outline the estimation strategy for LSTVAR model. The model is specified following the discussion on potential caveats and usual practice concerning the small open economies. To the best of our knowledge, no resembling analysis for CEE region has been pursued, which rules out a comparable specification. We include the vector of endogenous variables as above. Vector of exogenous variables is left out. Hence, the new identification scheme is modified

[11] Nigeria is the largest oil producer in Africa.

and reads:

$$\begin{bmatrix} \epsilon_t^1 \\ \epsilon_t^2 \\ \epsilon_t^3 \\ \epsilon_t^4 \\ \epsilon_t^5 \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ k_{21} & 1 & 0 & 0 & 0 \\ k_{31} & 0 & 1 & k_{34} & 0 \\ k_{41} & k_{42} & k_{43} & 1 & 0 \\ k_{51} & k_{52} & k_{53} & k_{54} & 1 \end{bmatrix} \begin{bmatrix} \mu_t^{ip} \\ \mu_t^{pr} \\ \mu_t^{ir} \\ \mu_t^{er} \\ \mu_t^{cr} \end{bmatrix}.$$

Again, non-policy block contains output and price level, implying that the monetary policy affects these variables only with a lag. Policy block consists of exchange rate and credit. Monetary authorities react to contemporaneous exchange rate fluctuations, but not to price level changes due to the inflation targeting horizon, which is between 12 and 18 months.

3.3.5 Dataset

Our sample is restricted to the 1998:M01-2012:M03 period, as the year 1998 coincides with the establishment of full-scale inflation targeting. Switch from fixed exchange regime may have changed the structural relationships among the variables, which could result in misleading conclusions. Industrial production (seasonally adjusted), money market rate (three-month) and monetary aggregate M2 are obtained from IMF International Financial Statistics (IFS).

CNB's official monetary policy instrument is the two-week repo rate; however, it is discontinuous and inconvenient for the purpose of our study. Instead, we use the three-month money market rate, which is the practice employed in mainstream studies on the Czech monetary transmission mechanism (Arnoštová & Hurník 2005; Borys *et al.* 2009; Havránek *et al.* 2010; Franta *et al.* 2012). Monetary aggregates do not appear as overly relevant for the conduct of the Czech monetary policy.

Price level is approximated by the net price index, which is available only internally at CNB. This measure excludes regulated prices. EUR/CZK nominal exchange rate and GDP are retrieved from the Eurostat database. Since the GDP series are available only at quarterly frequency, we interpolated them using the quadratic-match average procedure. Variable total credit to residents and nonresidents is obtained from CNB's public database ARAD, while the stock price index is downloaded from Prague Stock Exchange. All three exogenous variables – euro area GDP, price index and money market rate are retrieved from the Eurostat database. The exact description of the dataset is provided in Table A.1. Figure A.1 shows the logarithmized time series.

3.4 Results

This section presents the results and discusses their potential pitfalls. We first estimate the Linear VAR applying two different identification schemes. The series were further tested for the presence of threshold. In the case no threshold can be detected, Linear VAR is more efficient.

3.4.1 Linear Specification

Our first Linear VAR specification contains vectors of endogenous and exogenous variables as outlined in the Section 3.3.2. The model is estimated using detrended variables and the strategy outlined above. We choose the lag length of two, as suggested by the Final Prediction Error (FPE) and Hannan-Quinn (HQ) information criteria. Other criteria suggest higher lag order, which would be impractical in the model with eight variables and 171 observations. AR roots of characteristic polynomial lie inside the unit circle, indicating that the specification is stable.[12]

The results are presented in Figure A.3. We compute both orthogonalized and generalized impulse responses. As of the former, one Cholesky monetary policy innovation is associated with the fall in prices and credit and exchange rate appreciation. The response of industrial production is not clear: in the wake of monetary tightening, it appears to rise, declining after roughly seven months. Wide confidence bands indicate that this response is associated with a very high degree of uncertainty. Prices fall immediately, bottoming out around the fifteenth month. Exchange rate overshooting is evident: following monetary contraction, exchange rate appreciates (reaching the peak after 2-3 months), gradually depreciating afterwards. Credit declines as well, with the bottom between five and ten months.

Different identification scheme notwithstanding, our results roughly confirm previous findings on the Czech monetary transmission, most notably from Borys *et al.* (2009) and Havránek *et al.* (2010). The inclusion of a turbulent 2009-2012 period does not alter the impulse responses considerably. This is consistent with Franta *et al.* (2012), who find that the Czech output and prices are becoming increasingly responsive to monetary policy shocks. Such a result is attributed to anchored inflation expectations and financial deepening. In comparison to Arnoštová & Hurník (2005), no evidence of a price puzzle is re-

[12] The output is available upon request.

ported. This is probably due to the longer sample and the fact that we exclude pre-1998 period from the analysis.

Next, we compute the generalized impulse responses. GIRF circumvents the issue of identification and is therefore expected to provide more robust results. As stressed in the methodological section, when one variable is shocked, others are allowed to vary as implied by covariance. Mean impulse response function is calculated based on a certain number of simulations. In the context of our model, no substantial difference is reported: the responses to monetary tightening are in line with those previously reported. Some confidence bands are narrower, probably reflecting more realistic assumptions.

For more information on the interaction between various series, we perform the Granger causality test (Table A.3). The results are not interpreted as hints about the true causal relationship, but rather as the contribution to forecast of some particular variable. Credit seems to predict the interest and exchange rate, which is in line with our ordering scheme. Interest rate does not predict prices and industrial production in the current period, which is also consistent with the ordering. As far as the foreign variables are concerned, euro area GDP predicts industrial production at a 99% confidence level. The same can be stated about the euro area and domestic price level. No similar relationship is found between the foreign and domestic interest rate.

Variance decomposition is telling about the impact of foreign variables. We present the estimates in Figure A.4. Euro area GDP explains up to 40% of variation in domestic industrial production and 20% of variation in domestic price level. Another 20% of price fluctuations is explained by the euro area price level. Finally, 20% of credit variance is attributed to changes in the euro area GDP. Overall, foreign developments are fairly relevant for the Czech economy. This intuitive result is consistent with the findings reported in other relevant studies. Maćkowiak (2005) decomposes the sources of variation in aggregate output and prices for the three CEE countries – Czech Republic, Hungary and Poland, finding that 30% of output and 35% of price variation in the Czech Republic can be attributed to foreign shocks. Next, Jiménez-Rodriguez *et al.* (2010) address the impact of foreign shocks on key domestic variables in CEE economies, finding that positive commodity price shock has a significant positive effect on industrial production and price level in the Czech Republic.

Further, positive shock to foreign industrial production boosts the domestic industrial production, yielding a considerable exchange rate appreciation. Similar results are reported in Havránek *et al.* (2010).

The specification above produces estimates consistent with the theory and other studies. Due to the severe overparameterization, it is impractical in the context of LSTVAR. Hence, we estimate VAR(2) employing the identification scheme described Section 3.3.4 (Figure A.5) Impulse responses are somewhat counterintuitive: monetary shock appears to boost the industrial production, which declines slightly between the seventh and the tenth month and rises again. Prices fall immediately, but rise above zero already in the second month, declining again to bottom out around the fifteenth month. There is some evidence of an exchange rate puzzle; however, exchange rate appreciates soon after the monetary tightening. LSTVAR specification yields even more puzzling results: industrial production does not decrease over the response horizon. Similar situation is with the prices. Credit and exchange rate react roughly as described by the theory.[13] It is worth emphasizing that the impulse responses are fairly similar under the two regimes, implying the absence of a clear threshold effect. Clearly, these results are dubious and not in line with the relevant studies.

A number of explanations can be put forward. First, as outlined above and proven in different studies, foreign shocks do have a profound impact on the Czech economy. Their omission represents a serious drawback of our parsimonious specification. The results also deviate from what has been found about other small open economies. However, the studies in question have been conducted for the countries with different economic structure than the Czech Republic and employed somewhat different estimation strategies. For instance, Çatik & Martin (2012) and Bordon & Weber (2010) explore monetary transmission in the context of prior- and post-reform setting for Turkey and Armenia respectively. In both cases, the threshold is clearly detectable around the time of policy switch. On the other hand, we restrict our sample to post-1998 period, which corresponds to the adoption inflation targeting. Next, Saxegaard (2006) employs involuntary excess liquidity in the banking sector as a threshold variable. In the context of Subsaharan Africa, the justification is that the excess liquidity held in the banking sector bears inflationary potential, hence invalidating the effects of monetary policy. This would be unreasonable in our instance, as the Czech banking sector is subject to a set of macroprudential

[13] The output is available upon request.

rules and heavily supervised. Further, CNB has managed to anchor the inflation expectations, thereby gaining a high degree of credibility. Our specification is nearly identical to that in Jääskelä (2007) and Fuchun (2010); however, Australia and Canada are large economies of a different structure, which renders different results.

We attempt to solve this issue by augmenting the model with one additional variable. Degrees of freedom are being further restricted, albeit not substantially. Hence, we need to pick a variable exerting a strong influence over the Czech economy and/or excellent forecasting performance. World commodity price index is frequently used; however, it does not change our results considerably. Stock price index may meet these requirements, since they are believed to be determined by the macroeconomic fundamentals. As such, they are highly sensitive to macroeconomic and changes in future economic prospects. Moreover, in a small open economy, stock prices might reflect the prospect of external changes, which would later impinge on the real economic activity. Havránek *et al.* (2010) provide evidence of a consistently good forecasting ability of PX index. Our final identification scheme is therefore altered to include the stock market variable:

$$\begin{bmatrix} \epsilon_t^1 \\ \epsilon_t^2 \\ \epsilon_t^3 \\ \epsilon_t^4 \\ \epsilon_t^5 \\ \epsilon_t^6 \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 \\ k_{21} & 1 & 0 & 0 & 0 & 0 \\ k_{31} & 0 & 1 & k_{34} & 0 & 0 \\ k_{41} & k_{42} & k_{43} & 1 & 0 & 0 \\ k_{51} & k_{52} & k_{53} & k_{54} & 1 & 0 \\ k_{61} & k_{62} & k_{63} & k_{64} & k_{65} & 1 \end{bmatrix} \begin{bmatrix} \mu_t^{ip} \\ \mu_t^{pr} \\ \mu_t^{ir} \\ \mu_t^{er} \\ \mu_t^{px} \\ \mu_t^{cr} \end{bmatrix}.$$

The model is estimated on two lags, as suggested by the LR, AIC and SC information criteria (Table A.2). Engle-Granger coinegration test suggests the absence of long-term relationship among the variables. Stationarity is not explicitly tested for, as AR roots of characteristic polynomial lie inside the unit circle (Figure A.2). Impulse responses are presented in Figure 3.3. Industrial production slightly rises at the beginning, declining around the fifth month to bottom out between the tenth and the fifteenth. It subsequently rises and dies out by the end of forecasting horizon, which comprises 60 months. Again, the response over the first ten months is surrounded by a high degree of uncertainty. In comparison to the previous specifications, the response of industrial production is equally ambiguous. Notably, there is essentially no evidence of a price puzzle. Exchange rate appreciates, while the stock prices and credit fall in the wake of monetary tightening. Overall, these results corroborate the findings from the other studies.

3.4.2 Nonlinear Specification

Our next venture is to test for the presence of nonlinearities in the system above. If no nonlinearities are detected, LSTVAR is redundant. The number of lags is based on conventional lag length criteria, i.e. we apply the same lag as in baseline linear specification.

Pre-Testing for Nonlinearities

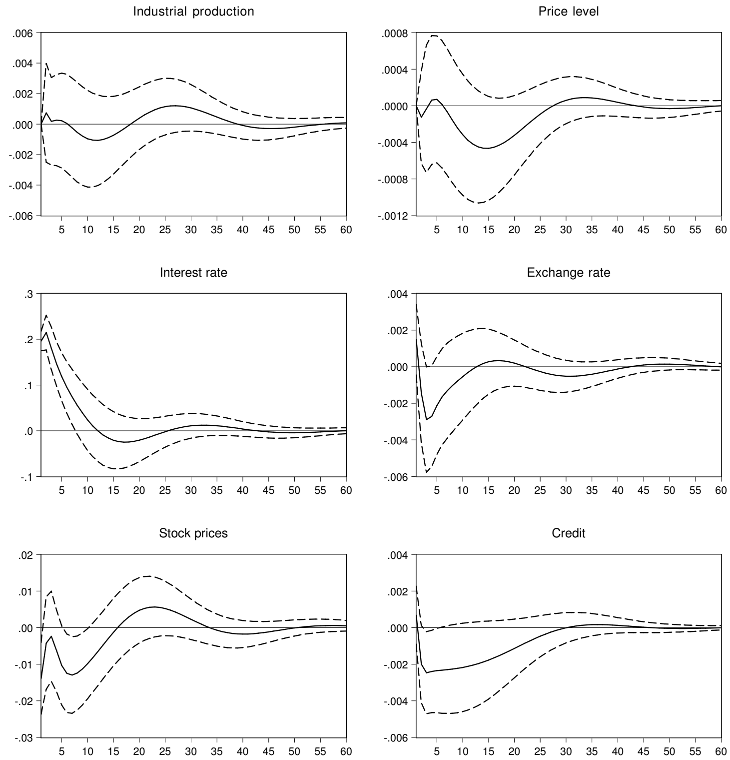
We test the null hypothesis of linear specification against the alternative. Since the probability distribution is different than χ^2, we apply the bootstrap procedure described in Weise (1999). Bootstrapped p-values are reported in the Table 3.1. The first five columns indicate test statistics for each equation separately at different lags of the transition variable, while the last one is related to the system as a whole.

Table 3.1: Lagrange Multiplier test for nonlinearity

Transition variable	**Dependent variable** ip	pr	ir	er	cr	LR
ip_{t-1}	0.9017	0.6935	0.9617	0.9632	0.9956	0.0009
ip_{t-2}	0.7742	0.9733	0.9763	0.9619	0.9892	0.0194
pr_{t-1}	0.4794	0.9131	0.8083	0.0922	0.5881	0.0141
pr_{t-2}	0.5633	0.9234	0.2596	0.0101	0.8521	0.0110
ir_{t-1}	0.9610	0.9629	0.9972	0.9996	0.6844	0.0000
ir_{t-2}	0.8516	0.7134	0.9992	1.0000	0.8483	0.0000
er_{t-1}	0.9989	0.9105	0.9996	0.9996	0.9999	0.0001
er_{t-2}	0.9994	0.9426	0.9999	0.9998	0.9997	0.0000
cr_{t-1}	0.7089	0.5811	0.9480	0.3469	0.7504	0.0001
cr_{t-2}	0.0694	0.5812	0.7243	0.3384	0.8642	0.0001

LM test: H_0: Linear equation vs. H_1: Nonlinear equation
LR test: H_0: Linear system vs. H_1: Nonlinear system

Figure 3.3: Impulse responses from Linear VAR, 95% confidence intervals



When the second lag of credit is used as the transition variable, the hypothesis of linearity is rejected for industrial production at 90% level. Intuitively, when the system is shocked, credit acts to induce the nonlinear response of industrial production. We also find evidence of nonlinearity in exchange rate equation when the first and the second lag of price level are used as the transition variables. LR test detects nonlinearity in the system for each of the transition variables considered. However, in all the other cases, it is impossible to identify in which particular equation the nonlinearity occurs. We acknowledge that this may be indicative of omitted variable bias, which we attempt to alleviate in further sections. For now, we inspect cr_{t-2}, pr_{t-1} and pr_{t-2} as potential transition variables.

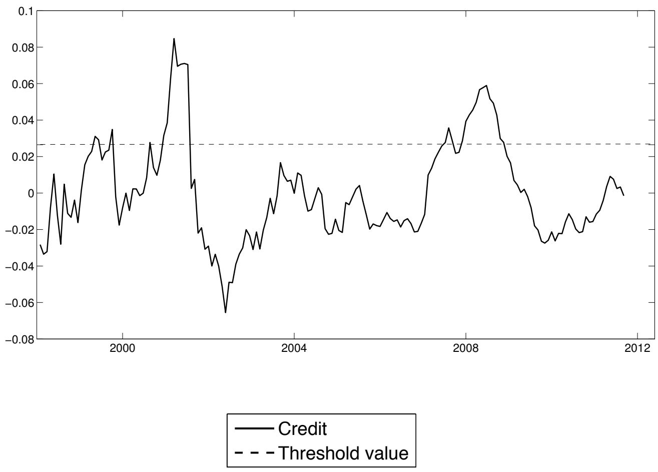
Optimal Threshold Value

The combination of optimal threshold and smoothness parameter value is selected using the grid search procedure described above. In the instance when cr_{t-2} is employed as the transition variable, the optimal combination of threshold and smoothness parameter is (0.0265, 1.4576). These values are significant at 99% level. When pr_{t-1} and pr_{t-2} are used as the switching variables, the optimal combination in both cases is (0.0068, 0.4775). However, neither of them is significant (p-values are 0.4775 and 0.3663 respectively). Conducting the grid search over entire range of the price variable, i.e. including the extreme 15% values is highly indicative. For both pr_{t-1} and pr_{t-2}, the optimal threshold value shifts to 0.0204. This observation is in the upper 3% extreme, which entails an important conclusion: price level threshold in the Czech Republic is too high to be of a practical importance. This finding should not be interpreted in a sense that the Czech economy behaves linearly. It should rather imply that the price level does not compromise the effectiveness of monetary policy, whereas the same may not be true about the credit growth. In line with the test results, our selection of the threshold variable pertains to cr_{t-2}.

As a way of judging the sensitivity of estimates, Weise (1999) suggests estimating the equation (3.15) employing the threshold specification ($\gamma \rightarrow \infty$). Threshold values remain unchanged in all instances.

We present the motion of credit variable with respect to threshold in Figure 3.4. Credit was over the threshold in periods 2001:M1-2001:M8 and 2007:M1-2009:M4. This finding is in line with the major developments in the Czech financial system. Banking sector was hit by a deep crisis during between 1997 and 1999, when the share of non-performing loans reached 30% (Frait *et al.* 2011). In a restructuring period, most of the banks were acquired by EU financial groups. Consequently, more loans were extended to the households and credit growth picked up sharply in 2001. Credit was also close to the threshold prior and immediately upon the EU accession, but did not cross it. Between 2001 and 2008, household borrowing displayed a high growth rate, which was not the case with the corporate sector. Further, a negligible share of FX loans did not expose banks to the exchange rate risk. Finally, the funding was generally provided on the basis of domestic deposits instead of foreign sources. All these factors acted in conjunction to prevent the scale of credit crunch comparable to that in some other CEE economies. Credit variable crossed the threshold in 2007:M1 and remained firmly above until 2009:M4. In line with the fact that no large-scale credit crunch took place, it fell below the threshold only in 2009:M5, with no profound deviations from the trend.

Figure 3.4: Transition variable path

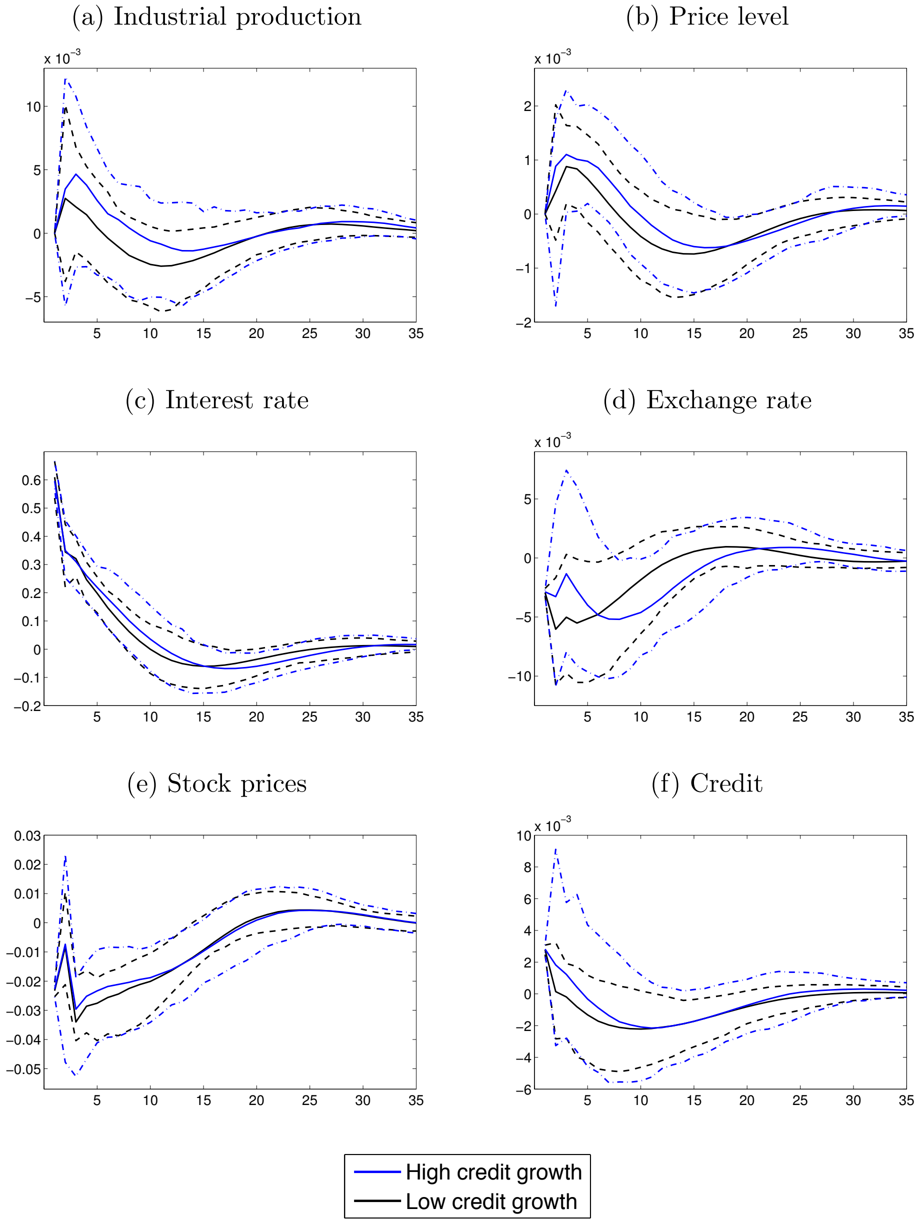


Estimation Output

We proceed with the LSTVAR estimation, which is suggested as optimal by the LM test and the detection of threshold value. This empirical exercise attempts to capture the asymmetries in the Czech monetary transmission. To test for the threshold effects, we compute the generalized impulse responses for two different scenarios, when the economy starts under and above the threshold. Regime switches are allowed throughout the duration of response, while the coefficients are changing accordingly. Responses from two different regimes, along with the corresponding 90% confidence bands are presented in Figure 3.5. To discern between the effects of one Cholesky monetary policy innovation, selected responses at different horizons are provided in the Table 3.2.

The sign of the responses computed from LSTVAR specification is somewhat counterintuitive. Monetary shock induces a rise in industrial production in both regimes. However, this should not be perceived as an anomaly of LSTVAR, considering the fact that the resembling response was evident in all previous Linear VAR specifications. It should be noted that the smallest puzzle occurred in the specification containing a block of foreign variables. This reflects the importance of external factors in the dynamics of the Czech economy. In our case, when the economy is initially below the threshold, industrial production declines in the fifth month, reaches the bottom in the eleventh and gradually dies out afterwards. This picture is different if the economy starts off above the threshold: it rises more than in the previous scenario and falls below zero only slightly around the fifteenth month. Negative response is deeper in

Figure 3.5: Impulse responses from LSTVAR, high vs. low credit growth, 90% confidence intervals



the former case, which indicates the presence of a threshold effect. Further, confidence bands are much wider in the latter case. Apart from more uncertainty under the high credit growth, this is also indicative about the lower number of observations pertaining to this particular regime. Unlike all previous specification, a large price puzzle emerges. The puzzle is evident in both regimes, albeit more under the high credit growth. Consequently, the decline in prices is relatively higher in a low growth scenario, which reflects the presence of a threshold effect. Importantly, the confidence bands are narrower than those of an industrial production. Large puzzle is clearly an anomaly, which we attempt to explain further.

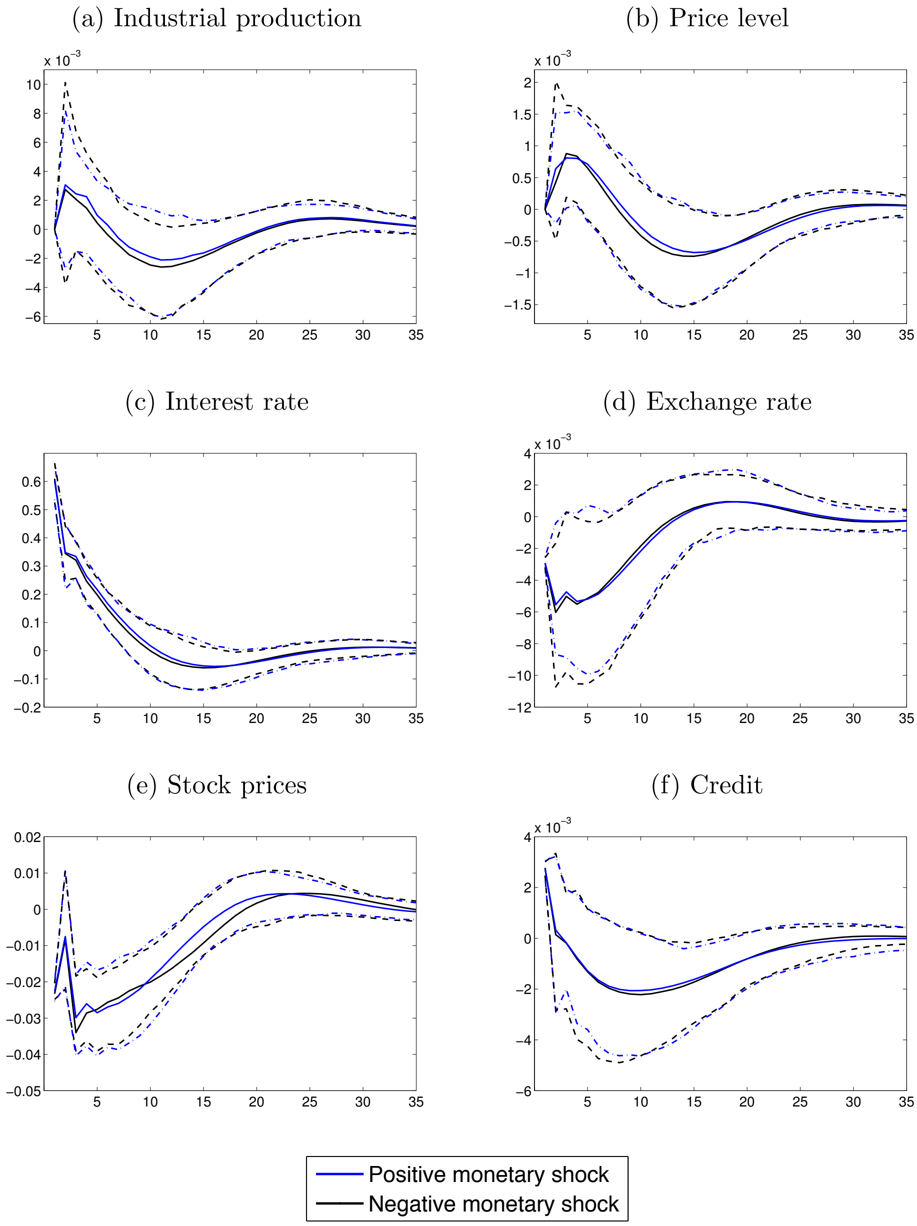
Table 3.2: Selected responses to +1 S.D. shock, high vs. low credit growth

	Industrial production		Price level		Exchange rate	
Horizon	**6 m.**	**12 m.**	**6 m.**	**12 m.**	**6 m.**	**12 m.**
Linear	-0.0001	-0.0012	0.0000	-0.0004	-0.0018	0.0000
Low growth	-0.0003	-0.0025	0.0004	-0.0006	-0.0047	-0.0005
High growth	0.0015	-0.0011	0.0008	-0.0003	-0.0048	-0.0033

The response of interest rate and stock prices to a monetary policy shock in the two regimes largely overlaps. This finding suggests that the reaction of monetary authorities to accelerated credit growth is not nonlinear. More specifically, if the credit growth passes a certain threshold, CNB does not raise the interest rate disproportionately. However, Vašíček (2011) suggests that the monetary policy rule under the conditions of financial distress is nonlinear. Resembling reaction of stock prices to a monetary tightening in the two regimes indicates that the investors do not account for possible threshold effects. Namely, monetary tightening is expected to have uniform effects on the economy, regardless of the degree of credit growth.

As far as the exchange rate is concerned, appreciation is evident under both credit regimes. In a high growth regime, it hovers close to zero in the first four months, subsequently declining and bottoming out around the ninth month. However, the confidence bands are relatively wide. This scenario is somewhat different in a low regime, where the exchange rate firmly appreciates in th first few months. There is no considerable difference in the maximum degree of appreciation. Finally, the response of credit is above zero in the first three months. Afterwards, it decreases and dies out in about two years. This reaction is deeper of the economy is initially below the threshold.

Figure 3.6: Impulse responses from LSTVAR, positive vs. negative shock, 90% confidence intervals



Next, we test for the relative effects of expansionary and contractionary monetary shocks. In both instances, the economy starts off below the threshold. The responses to expansionary monetary shock are rescaled by (-1) to ensure comparability. Figure 3.6 indicates no significant differences in the effects of positive and negative shock. The impulse responses of all variables in the two scenarios largely coincide, along with the confidence bands. Timing of the responses is similar as previously. Industrial production initially rises, falling below zero in the seventh month and bottoming out in the twelfth. Price puzzle emerges, with the prices reaching the bottom around the fifteenth month. Strong exchange rate appreciation is evident in the first six months, followed by gradual depreciation. Stock prices remain firmly below zero throughout the duration of the response. Finally, credit is positive in the first three months, gradually declining afterwards.

Further Discussion

Comparison of the results from Linear VAR and LSTVAR is indicative about the peculiarities of generalized impulse responses. First, the fall in industrial production in Linear VAR is roughly between the two impulse responses from LSTVAR. However, the initial rise in LSTVAR is considerably higher. The same can be stated about the price level. Importantly, price puzzle from Linear VAR is almost negligible. Confidence bands of industrial production response are much wider in LSTVAR, whereas those of price level are comparable. Stock prices display somewhat stronger responses in LSTVAR. Exchange rate and credit react similarly across the two specifications. Further, the reaction of an interest rate is stronger in nonlinear model. It is worth noting that in all instances, impulse responses become insignificant after roughly 30 months.

Interpretation of the results warrants further discussion on the essence of generalized impulse responses. The first simulation is performed by drawing a random vector of shocks from regression residuals and propagating them through the system. In the second simulation, the same vector is used, plus additional perturbation which is constructed from structural shock to the selected variable using the recursiveness assumption. The difference between these two simulations represents generalized impulse response (Mandler 2010). This procedure is repeated for each regime using 1000 draws of random shocks. Hence, the interaction of shocks throughout the duration of response is allowed, which may magnify the effects of an initial one. Considering that the initial responses of industrial production and price level in Linear VAR are slightly above/close to zero, the simulation procedure described above might have magnified them, producing anomalies in LSTVAR. Similarly, this may be the reason behind somewhat deeper reaction of stock prices in LSTVAR.

We further draw a comparison between our results and other studies employing some version of threshold specification. In the study by Çatik & Martin (2012), the link between interest rate and prices in Turkey appears to be broken in a pre-reform period (prior to 2004). Its disappearance in the ensuing period is explained by the less accommodative monetary policy. However, it is worth reiterating that this study uses a long data sample and three exogenous variables, which may result in more reliable estimates. In the post-reform period, monetary tightening is followed by the initial output increase. The authors justify this by a shorter sample period and therefore, less precise finding. This claim very well applies to our analysis. Nearly all the responses die out around the twenty-fifth month, which resembles our results. In a similar vein, Bordon & Weber (2010) document the price puzzle on Armenian sample spanning between 2006:M1 and 2010:M5. An abundant evidence of price puzzle is reported in Bathaluddin *et al.* (2012) as well. In spite of relatively long samples, studies on UK and US are no exception in terms of reported anomalies. Atanasova (2003) documents persistent rise in UK inflation three months upon the monetary contraction. Karamé & Olmedo (2002) report weakly functioning monetary transmission in US during recessions. A more recent study by Mandler (2010) lends support to this finding. However, one should not dismiss the fact that all of them employ relatively parsimonious specifications.

Our finding of symmetry related to the direction of shock corroborates other studies. The concept of sign-related asymmetries is theoretically appealing; however, the supporting empirical evidence arises only under the certain conditions. Atanasova (2003), Weise (1999) and Mandler (2010) find sparse evidence of asymmetries in the effects of positive and negative shocks. In TVAR specifications by Balke (2001) and Calza & Sousa (2005), asymmetries disappear once the nonlinear impulse responses are allowed for. Asymmetries have been detected in the studies by De Long & Summers (1988), Cover (1988) and Thoma (1994). However, if one controls for policy shifts, no evidence of asymmetries can be established (Weise 1999; Ravn & Sola 1999). Our analysis satisfies both conditions: we allow for nonlinear responses and include the post-1998 period only. Sensibly enough, positive and negative shocks have symmetric effects.

Several issues deserve further discussion. First, there is an obvious trade-off between the parsimonious specification and the efficiency of estimates. This is evident in the context of both large and small open economies. The issue becomes more pronounced in the latter case, as the small economies typically require the inclusion of the variables capturing external developments. Jarociński (2008) notes that, due to short sample periods, the existing VAR studies for post-communist countries are sensitive to small changes in specification and suffer from a high degree of uncertainty. To alleviate this problem, he suggests pooling the available information on different countries. We attempt to conduct such an analysis on a panel of CEE economies in the following chapter.

Next, VAR analysis is consistently plagued by the price puzzle. Employing a large meta-analysis, Rusnák *et al.* (2011) find several possible explanations of this phenomenon. First, some publication bias is evident against the studies reporting the long run puzzle. In a short run, there is a profound uncertainty regarding the duration of monetary transmission itself. Moreover, cost channel theory implies that the firms' optimal response to a monetary tightening are higher prices, due to their dependence on credit. Hence, there is some theoretical justification for an initial rise in prices. On the other hand, price puzzle is predominantly related to model misspecification - omission of the commodity price index, output gap or the use recursive identification scheme.

Finally, Égert & MacDonald (2008) underline potential dysfunctionality of the broad credit channel in CEE countries. The practice of raising the funds publicly is severely underdeveloped, which gives banking sector a certain monopoly. Hence, higher degree of financial distress could act as a powerful propagator of monetary shocks. This is not the case in reality, as the banks and other financial corporations can circumvent domestic constraints by borrowing from their mother companies. On the other hand, parent institutions are subject to the credit channel in euro area, which is still insufficient to render it operative in CEE region.

3.5 Sensitivity Analysis

To check for the robustness of our findings, we employ two different specifications. In the first one, industrial production is replaced with GDP. Quarterly series were converted to a monthly frequency using the quadratic-match procedure. Number of lags and identification scheme remain the same as above. The results are presented in Figure A.6 and Figure A.7. Our results remain robust to the specification including GDP. The responses of variables to all the shocks largely coincide to those in the previous specification. Monetary policy shock has a negative impact on GDP, price level, stock prices and credit, while the exchange rate appreciation is evident. This impact is stronger under the conditions of restricted credit. On the other hand, impulse responses to positive and negative shocks are largely symmetric.

Our second robustness check employs different detrending technique. Specifically, we use the HP filter but for the time frequency of 60 months. This duration of the Czech business cycle is in line with Poměnková & Maršálek (2011). Further, different identification scheme is employed: exchange rate is ordered after the policy rate, implying that monetary authorities do not affect it contemporaneously. The results presented in Figure A.8 and Figure A.9 largely correspond to our initial results: monetary policy loses its clout in a high credit growth regime, whereas the effects of positive and negative shocks display no asymmetries.

Chapter 4

Financial Structure and Monetary Transmission in CEE Region

This chapter is an attempt to model monetary transmission asymmetries for CEE region. As emphasized previously, short sample periods plague individual country findings by a high degree of uncertainty. Further, the use of total credit as a sole measure of financial conditions may be deficient. Henceforth, we attempt to remedy for the flaws of one-country analysis. Using the sample of eight CEE states – Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Slovakia and Slovenia, we quantify the contribution of financial structure to the overall effectiveness of monetary policy. To capture macrofinancial linkages in a more comprehensive manner, we construct two indicators of financial structure – credit dependence and banking sector competition.

4.1 Stylized Facts

Banking sector in CEE region has been changing rapidly since mid-1990s. Three crucial evolutionary factors included structural reforms, credit boom and post-2008 credit crunch. This section presents a short theoretical background and some of the pertinent stylized facts, which place our further analysis in the context.

Theoretical literature tends to ascribe credit booms to financial deepening or financial accelerator mechanism. Specifically, economic growth is believed to create a higher degree of financial intermediation. The exact causal relationship is unclear, however, as more financial services may contribute to the economic growth. As of the financial accelerator mechanisms, credit expansion emanates from over-optimism about future earnings, which increases the collateral value and consequently, borrowing capacity.

Over the past decade, CEE region underwent a period of rapid credit growth. The ramifications of a consequent downturn were profound, but divergent across the countries. Acknowledging peculiarities of individual credit developments, a few region-wide tendencies can be identified. First, credit growth picked up around 2000, which coincides with the end of major structural reforms and more robust economic growth. Second, the severity of credit crunch differed based on the soundness of banking sector and FX exposure. We analyze these facts below.

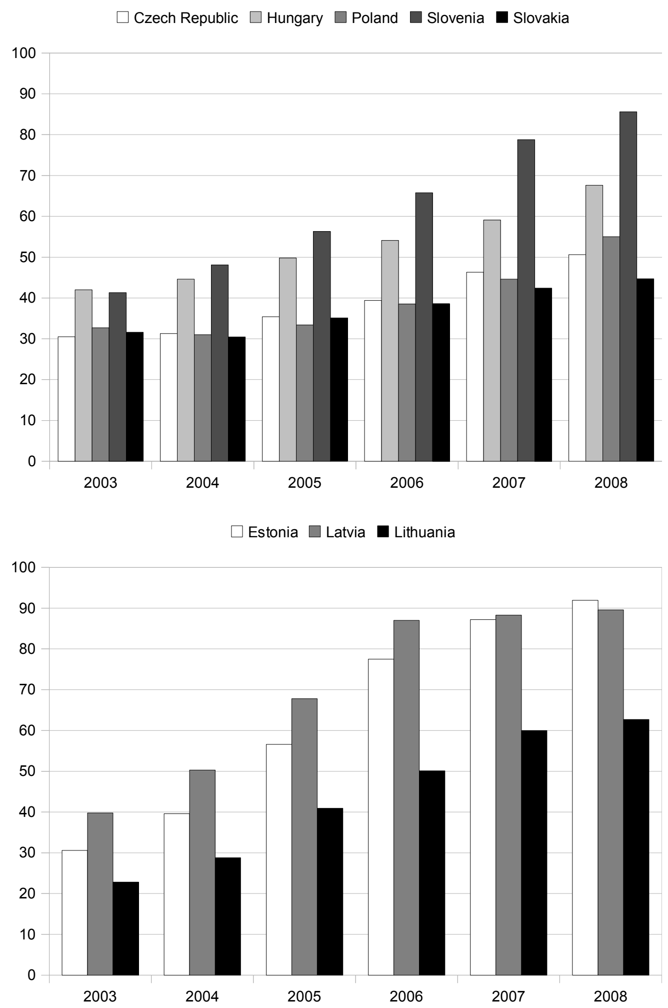
Financial system in CEE economies is predominantly bank-based, with the banks owning roughly 85% of total financial sector assets (Égert *et al.* 2006). Another important observation is that the vast majority of banks is foreign-owned, whereas the share of state-controlled entities is low to negligible.[1] Capital markets remain underdeveloped as compared to the euro zone. Hence, credit is the major source of financing. Private credit-to-GDP ratio in Figure 4.1 reveals a more nuanced picture. Between 2003 and 2008, credit growth reached very high levels in Estonia and Latvia. Slovenia and Hungary exhibited similar pace, albeit this ratio remained somewhat lower. Credit expansion remained subdued in the Czech Republic, Poland and Slovakia, with some acceleration visible in 2007-2008.

Dissecting the total private credit into corporate and household loans reveals that the corporates were equally avid borrowers (Figure 4.2). The average corporate credit-to-GDP ratio ranged between 14% in Poland to 50% in Slovenia. Household lending was predominantly mortgage-based in the Baltics, the Czech Republic and Hungary. On the other hand, Poland, Slovenia and Slovakia extended more consumption-related loans.

The share of FX loans in some economies was substantial (Figure 4.3). This phenomenon was pronounced in the Baltic states and Hungary, whose monetary frameworks rested on exchange rate targeting or currency board regimes. On the other side of the spectrum, FX loans remained negligible in the Czech Republic. Another important development pertained to the cross-border lending. As an outcome of financial integration, a number of countries encountered high portions of direct foreign lending, which is visible from the Figure 4.4. Again, the Czech Republic, Poland and Slovakia stand out as the exceptions. Foreign currency borrowing was the most accelerated in a mortgage segment,

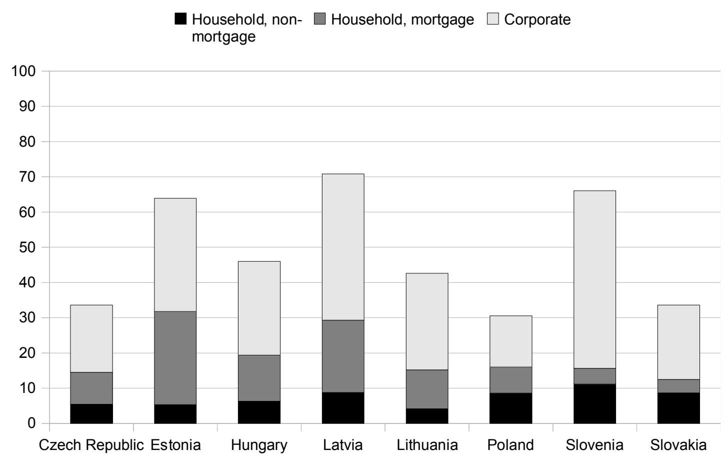
[1] As an exception, the largest Polish bank (PKO BP) is majority state-owned.

Figure 4.1: Domestic credit to private sector (in % of GDP)



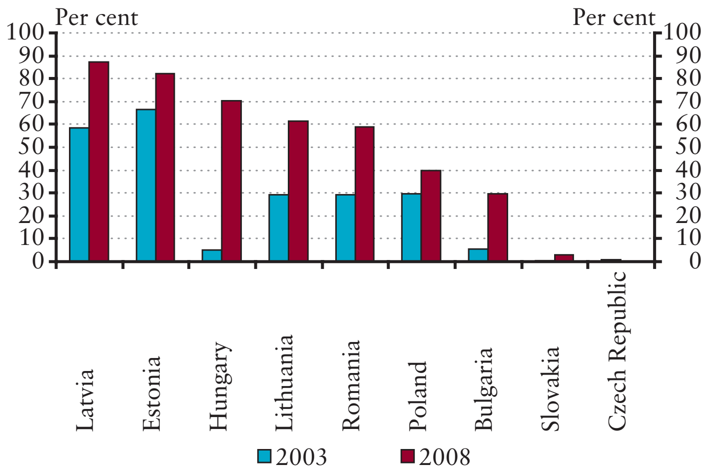
Source: EBRD (2009) and World Bank

Figure 4.2: Household vs. corporate credit, 2003-2008 averages



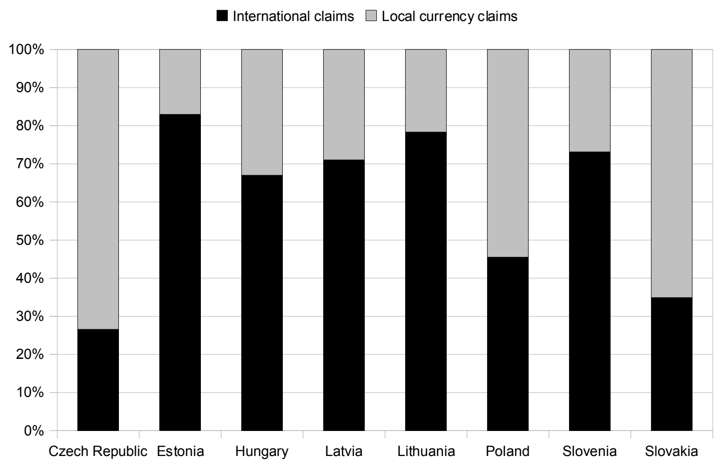
Source: EBRD (2009)

Figure 4.3: Share of FX loans within the banks' household loan portfolio, 2003 vs. 2008



Source: MNB (2009)

Figure 4.4: International vs. local currency claims as a share of total foreign claims, 2003-2008 averages



Source: BIS

due to the fact that long-term rates remained prohibitively high in some of the countries. In the context of our analysis, it is worth emphasizing that cross-border lending dampens the effectiveness of monetary policy.

Credit crunch which unfolded after 2008 was unprecedented in scale. First, as a result of global downturn, parent institutions faced liquidity shortages, which induced lending cuts in the region. Second, foreign currency lending stumbled due to the fall in demand and exchange rate depreciation. Baltic states and Hungary were the most adversely affected, in line with the peculiar credit developments. Some turbulence notwithstanding, the Czech Republic, Poland and Slovakia displayed a remarkable degree of resilience.

Market structure is an important determinant of banking sector trends. Banking sector competition indicators (Figure 4.5) reveal rather different cross-country developments. Herfindahl-Hirschman index indicates that Estonian and Lithuanian banking sectors are strongly concentrated.[2] In these countries, five largest banks hold 95% and 85% of total bank assets respectively (Economic and Financial Affairs Directorate General 2010). Slovenia is above the average, whereas Hungary and Poland remain below. Latvia, Slovakia and the Czech Republic display similar trends. High net interest margin confirms weak competitive pressures in the Estonian market.[3] Lower concentration notwithstanding, the margin remained firmly above the average in Hungary, probably reflecting higher premiums from 2006 onwards. Poland is also somewhat above the other economies, although with a decline evident from 2006. No large-scale differences across the remaining countries can be discerned.

4.2 Literature Survey

This section expands evidence on the conjunction between financial structure and monetary transmission. Following the discussion in the Section 2.1.1, we survey the literature on interest rate channel in CEE region. We emphasize that very few studies compare the actual strength of transmission channels across the region, which hinders our interpretation of results.

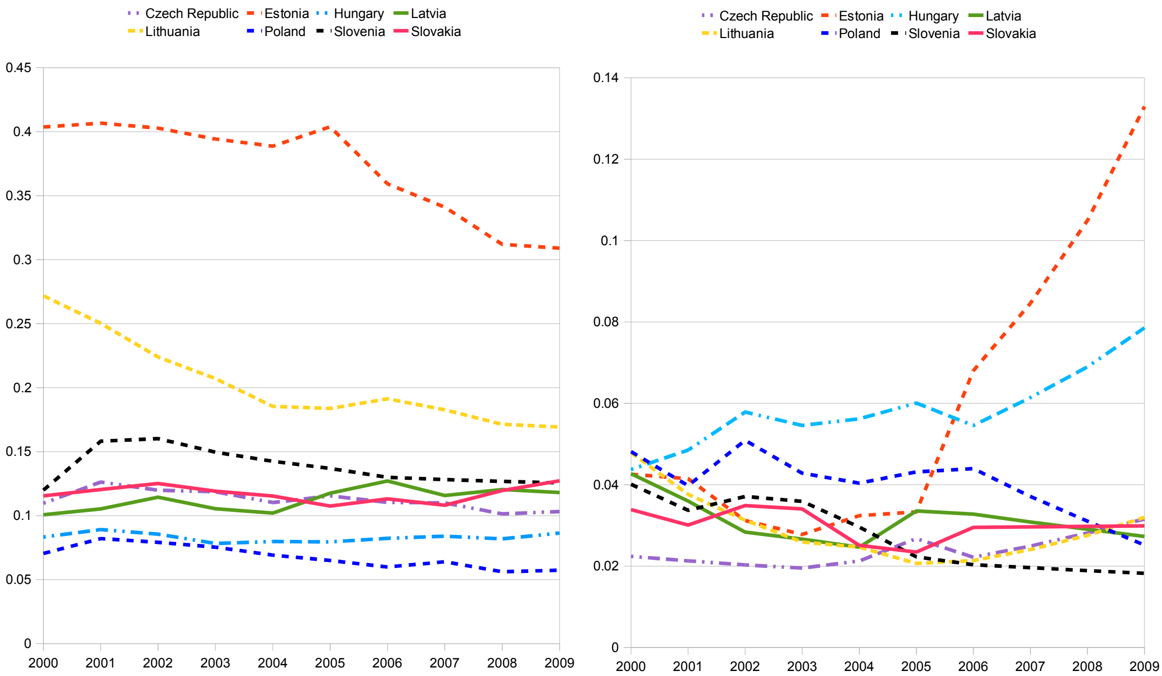
In one of the early studies, Kot (2004) establishes empirical relationship between the interest rate pass-through and the degree of banking sector com-

[2] Herfindahl-Hirschman index measures the banks' asset size in relation to an entire industry.

[3] Net interest margin is defined as the difference between lending and borrowing interest rate, divided by interest-bearing assets.

Figure 4.5: Banking sector competition indicators

(a) Herfindahl-Hirschman index (b) Net interest margin



Source: (a) ECB, (b) World Bank Financial Structure Database

petition for the Czech Republic, Hungary and Poland. Transmission in all three countries is regarded as sluggish and incomplete. Higher degree of competition may not enhance the interest rate pass-through, considering the fact that Poland and Hungary exhibit more competitive banking sector and weaker pass-through than the euro zone. Égert *et al.* (2007) compare the interest rate pass-through in CEE-5 and selected euro-area economies. The repercussions of monetary policy on long-term market rates are found to be weak, due to a highly volatile yield curve. Overnight rates are generally insulated from monetary policy shifts, while the pass-through strengthens for long-term deposit rates. Further, corporate lending rates are less rigid than the household rates. Individual country pictures are more nuanced. Czech Republic, Slovakia and Hungary exhibit little or no pass-through to household lending rates. Mortgage-related rates are versatile in Slovenia and Spain. Finally, in the case of Germany and Poland, the transmission to deposit rates is more pronounced than for lending rates. In a more elaborate study, Égert & MacDonald (2008) summarize all available findings on CEE region and calculate the average pass-through to a range of rates. The strongest pass-through is detected for short- and long-term corporate lending rates, while that to the consumer loans is weak. Higher responsiveness of corporate rates may emanate from the tighter competition in the sector of corporate banking. This is in line with Demel & Sikimic (2008), who document a highly competitive market for small- and medium-enterprise loans. As of specific countries, the pass-through is high in the Baltics and low in the Czech Republic and Slovakia. The results for Hungary and Poland are less clear-cut. It is worth highlighting that in CEE-5, the strength of transmission gradually diminishes.

The importance of credit in the economy can be examined through the functioning of monetary transmission channels. Strong amplification of monetary shocks through the components of credit channel indicates high relevance of the bank credit. Studies addressing this issue for CEE region are numerous. General support for the operativeness of credit channel notwithstanding, the scope of these papers does not allow us to draw overarching conclusions about each specific country.

Matousek & Sarantis (2009) test for the effectiveness of bank lending channel on a sample of eight CEE economies for the 1994-2003 period. Bank lending channel is found to be functioning in all the states, albeit considerable cross-country differences can be detected. Size and liquidity appear to play an important role in banks' response to a monetary policy shock. Aggregate loan supply is found to have a positive and significant effect on the real economic activity, which is telling about the importance of bank credit in CEE region. Elbourne & De Haan (2006) expand the aforementioned sample with Romania and Bulgaria to examine the conjunction between financial structure and monetary transmission. Employing three broad indicators of financial structure - indicators for the importance of small banks, health of the banking system and availability of alternative sources of financing, the study finds correlation only between the second indicator and the strength of monetary transmission. No systematic relationship of monetary transmission with banks' size and competition from other forms of finance is detected.

Summarizing evidence on individual countries is tedious, partly due to different time spans and a wide variety of econometric techniques employed. Early evidence is plagued by a high degree of uncertainty. Wróbel & Pawłowska (2002) find that the credit channel in Poland is generally operative, while the interest rate pass-through is comparable to that of the euro zone. Further, it is demonstrated that higher concentration in the banking market entails slower pass-through for the lending rates, but increases it for the deposit rates. Matousek & Sarantis (2009) pinpoint capitalization as a determinant of bank lending channel in Poland. This bolsters the finding of Wróbel & Pawłowska (2002), who argue that monetary contraction primarily affects poorly capitalized banks. In a more recent study, Łyziak *et al.* (2008) contest the previous results by delivering mixed evidence on the effectiveness of credit channel. Firms' balance sheet amplifies the effects of monetary shocks, but only weakly. Loan supply is almost unresponsive to monetary policy shifts, which is explained by the behavior dubbed as "buffer-stock". Namely, banks hold a high amount of liquid assets, which allows them to reduce the supply of short-term loans and insulate the portfolio of investment loans.

Horváth *et al.* (2004) examine the functioning of interest rate channel in Hungary, finding that the speed and completeness of pass-through to corporate deposits and loans are considerably higher than to the household instruments. As of the former, faster short-term and more complete long-term adjustment is evident. Specifically, 70% of an interest rate shock is transmitted in the first period of adjustment. This finding is attributed to a more intense competition in the corporate banking sector. In spite of this, corporate loan rates exhibit downward rigidity. Concerning the credit channel, Horváth *et al.* (2006) report that bank size, liquidity, capitalization and foreign ownership do play a role in the transmission of monetary shocks, albeit some of the findings are not robust across different specifications. Foreign-currency denominated corporate loans are fairly unresponsive to monetary policy changes. This contrasts to Matousek & Sarantis (2009), who document sparse evidence on the bank lending channel in Hungary.

Concerning the Czech Republic, the study by Fidrmuc *et al.* (2008) offers a degree of support for the operativeness of the balance sheet channel. In one of the earliest studies on the interplay between monetary policy and banking sector, Pruteanu (2004) argues that the Czech bank lending channel was weakly operative between 1999 and 2001, which coincides with the period of banking sector restructuring. Contrary to these findings, Matousek & Sarantis (2009) find that size, capitalization and liquidity are significant in all periods. However, their relative importance is difficult to asses. No unanimous conclusions about Slovakia can be drawn, since their distinctive bank features do not appear robust as factors of monetary transmission across different specifications. Ahtik (2012) finds evidence of a functioning bank lending channel in Slovenia. More profitable, liquid and deposit-funded banks are less responsive to policy changes. The impact of size is uncertain, whereas that of capitalization is not robust across specifications. In the context of our analysis, subsidiaries of foreign banks tend to be somewhat isolated from the impact of monetary policy. Finally, higher concentration is found to debilitate the efforts by monetary authorities.

The evidence on the credit channel in the Baltics is mixed. Köhler *et al.* (2005) find that small and well-capitalized banks react more intensely to a restrictive monetary policy. However, liquidity and capitalization appear to have counter-balancing effects, which renders the final impact of monetary policy indeterminate. In this context, Égert & MacDonald (2008) argue that the bank-level capitalization may not matter due to the fact that leverage of an entire system exerts the influence over the process of monetary transmission. The relevance of bank lending channel vanishes under the conditions of abundant foreign lending: in Köhler *et al.* (2005), euro money market rate turned out to be a highly significant determinant of bank lending behavior, which was not the case with domestic money market rate. Hence, high share of FX loans retards the effectiveness of monetary policy. Beņkovskis (2008) concludes on a similar note: albeit some evidence on the bank lending channel is present, FX loans are mostly unresponsive to policy changes. However, Juks (2004) demonstrates that this channel was functional prior to 2003, which coincides with the onset of excessive rise in foreign lending.

4.3 Methodology

Panel estimation techniques are tailored to provide a richer knowledge about cross-sectional and time-specific dynamics. With this additional information, one is able to obtain more efficient parameter estimates. Panels are used more avidly in microeconomics, typically by pooling a large number of cross-sectional units over a short period of time. Conventional estimation techniques are thus based on the properties of microeconomic panels. This is somewhat inappropriate in the context of macroeconomic panels, where the number of cross-sectional units tends to be small, while the time span is somewhat longer. This section is designed to put forward the major issues related to the macroeconomic panels and present Panel Conditionally Homogenous Vector Autoregression model (PCHVAR), which is employed as an estimation tool. For the sake of brevity, we do not discuss the problems related to the consistency of parameters. Readers interested in this topic should consult the standard textbook by Canova (2007) or Smith & Fuertes (2010).

Table 4.1: Literature overview

Study	**Country**	**Issue**	**Findings**
Kot (2004)	CEE-3 vs. euro zone	Transmission and bank competition	Weaker pass-through in CEE-3
Elbourne & DeHaan (2006)	CEE	Financial structure and monetary transmission	Weaker banks enhance the transmission
Égert *et al.* (2007)	CEE-5	Interest rate channel	Weak in the Czech Republic, Slovakia and Hungary; strong for mortgages in Slovenia
Égert & MacDonald (2008)	CEE	Interest rate channel	Weak in the Czech Republic, Slovakia and Hungary; strong in the Baltics; ambiguous in Poland and Slovenia
Matousek & Sarantis (2009)	CEE	Bank lending channel	Generally operative
Wróbel & Pawłowska	Poland	Credit channel	Generally operative
Łyziak *et al.* (2008)	Poland	Credit channel	Weakly operative, mainly due to a high liquidity of the banks
Horváth *et al.* (2004)	Hungary	Interest rate channel	Downward rigid corporate rates, otherwise operative
Horváth *et al.* (2006)	Hungary	Bank lending channel	Generally operative, but FX loans almost unresponsive
Fidrmuc *et al.* (2008)	Czech Republic	Balance sheet channel	Generally operative

Study	Country	Issue	Findings
Pruteanu (2004)	Czech Republic	Bank lending channel	Dysfunctional between 1999 and 2001
Ahtik (2012)	Slovenia	Bank lending channel	Operative, liquidity, funding sources and concentration matter
Köhler *et al.* (2005)	Baltics	Bank lending channel	Almost non-functional, due to a high share of FX loans
Beņkovskis (2008)	Latvia	Bank lending	Relatively dysfunctional
Juks (2004)	Estonia	Bank lending channel	Generally functional up to to 2003

4.3.1 Introduction to Dynamic Macro Panels

Consider the panel consisted of $i = 1, ..., N$ cross-sectional, $t = 0, ..., T$ time units and j lags. Generic equation capturing this setting takes the following form:

$$y_{it} = \sum_{j=1}^{p} A_j y_{i,t-j} + \epsilon_{it}, \qquad \epsilon_{it} \overset{i.i.d.}{\sim} (0, \Sigma_\epsilon), \tag{4.1}$$

where y_{it} is a $K \times 1$ vector of endogenous variables and A_j is a $K \times K$ matrix of coefficients. The key assumption is convergent cross-sectional dynamics, which entails homogenous slope coefficients. Individual-specific effects are usually captured by the time-invariant fixed or random effects. This estimation procedure is common in microeconomic research (Canova 2007).

The aforementioned restriction does not hold in a macroeconomic setting. Cross-sectional units are typically subject to different policies, whose ramifications need to be evaluated. Resorting to fixed effects in this instance is problematic. Consider the equation

$$y_{it} = \sum_{j=1}^{p} A_j y_{i,t-j} + A_{p+1}\rho_i + \epsilon_{it}, \qquad \epsilon_{it} \overset{i.i.d.}{\sim} (0, \Sigma_\epsilon), \tag{4.2}$$

where ρ_i represents the unobservable (unit-specific) effects. Allowing for heterogenous intercepts ρ_i while the slope homogeneity is still present yields serial correlation among the residuals, which results in inconsistent estimates (Sá *et al.* 2011). Heterogeneity bias may be resolved employing a mean group estimator, which estimates the time- and unit-specific parameters:

$$\epsilon_{it} = \mu_i + \lambda_t + v_{it}. \tag{4.3}$$

Here, μ_i represents unit-specific fixed parameter, λ_t stands for time-specific fixed parameter, while v_{it} is the remainder effect. Depending on the magnitude of v_{it}, dynamics may remain heterogenous even upon controlling for time- and unit-specific effects. Another problem is that heterogeneity v_{it} is treated as a nuisance. Thus, the mean group estimator does not lend itself to an individual country analysis. Clearly, we need a specification which allows to make inferences about the sources of heterogeneity, simultaneously exploiting the poolability of the data.

4.3.2 The Model

Following Georgiadis (2012a), we set up the Panel Conditionally Homogenous Vector Autoregression model (PCHVAR), which allows for heterogenous cross-sectional dynamics:

$$y_{it} = \sum_{j=1}^{p} A_j(z_{it}) y_{i,t-j} + \epsilon_{it}, \qquad \epsilon_{it} \overset{i.i.d.}{\sim} (0, \Sigma_\epsilon), \tag{4.4}$$

where $i = 1, 2, ..., N$ denotes cross-sectional and $t = 1, 2, ..., T$ time units, y_{it} is a $K \times 1$ vector of endogenous variables, z_{it} is a $M \times 1$ matrix of exogenous variables, ϵ_{it} is a vector of serially uncorrelated reduced-form disturbances, while A_j represents the $K \times M$ coefficient matrix.

The model above is the generalized version of Panel VAR from equation (4.1). Unlike the regular Panel VAR, the specification in (4.4) allows for different cross-sectional dynamics, but only to the extent of divergences in each unit's observed heterogeneities. To restate, the time series are pooled, but impulse responses differ based on the values of exogenous conditioning variables

z_{it}. Hence, this specification is indicative about the role of individual structural characteristics in observed heterogeneities. Similarly as the transition variable in LSTVAR, z_{it} is an essential determinant of model's dynamics.

The solution and the corresponding impulse response computation closely follow Georgiadis (2012a). Let $a_{j,sm}(z_{it})$ be the scalar coefficients from $A_j(\cdot)$, where $j = 1, 2, ..., p$, $s = 1, 2, ..., K$ and $m = 1, 2, ..., K$. Assume that each scalar coefficient can be approximated by a scalar polynomial in the conditioning variable z_{it}:

$$a_{j,sm}(z_{it}) \approx \pi(z_{it})\gamma_{j,sm}, \tag{4.5}$$

or, in the case of several conditioning variables, by a multivariate polynomial $\pi(z_{it})$, which can be obtained as a Kronecker product of two univariate polynomials in the conditioning variable:

$$\pi(z_{it})\gamma_{j,sm} = [\pi_1(z_{1i,t-1}) \otimes \pi_2(z_{2i,t-1})]\gamma_{j,sm}, \tag{4.6}$$

$$\pi_j(z_{ji,t-1}) = [\pi_j^{(0)}(z_{ji,t-1}), \pi_j^{(1)}(z_{ji,t-1}), ..., \pi_{j)}^{(\tau j)}(z_{ji,t-1})], \qquad j = 1, 2 \tag{4.7}$$

where $\pi(z_{it}) = [\pi_1(z_{it}), \pi_2(z_{it}), ..., \pi_\tau(z_{it})]$ is a $1 \times \tau$ vector with polynomials in z_{it} and $\gamma_{j,sm} = (\gamma_{j,sm1}, \gamma_{j,sm2,...,\gamma_{j,sm\tau}})'$ is a $\tau \times 1$ vector of polynomial coefficients.[4] $A_j(\cdot)$ can be rewritten as follows:

$$A_j(z_{it}) = \begin{bmatrix} \pi(z_{it})\gamma_{j,11} & \dots & \pi(z_{it})\gamma_{j,1K} \\ \vdots & \ddots & \vdots \\ \pi(z_{it})\gamma_{j,K1} & \dots & \pi(z_{it})\gamma_{j,KK} \end{bmatrix}$$

$$= \begin{bmatrix} \gamma'_{j,11} & \gamma'_{j,12} & \cdots & \gamma'_{j,1K} \\ \vdots & \ddots & \vdots & \\ \gamma'_{j,K1} & \gamma'_{j,K1} & \cdots & \gamma'_{j,KK} \end{bmatrix} \cdot \Big[I_K \otimes \pi'(z_{it})\Big] =$$

$$= \Gamma \cdot [I_K \otimes \pi'(z_{it})]. \tag{4.8}$$

[4] We may try to preserve the degrees of freedom by dropping the interaction terms between conditioning variables and obtaining the following approximations: $\tilde{\pi}(z_{it}) \cdot \gamma_{j,sm} = [1, \tilde{\pi}_1(z_{1i,t-1}), \tilde{\pi}_2(z_{2i,t-1})] \cdot \gamma_{j,sm}$ and $\tilde{\pi}(z_{ji,t-1}) = [\pi_j^{(1)}(z_{ji,t-1}), ..., \pi_j^{(\tau_j)}(z_{ji,t-1})], \quad j = 1, 2.$

The model in the equation (4.4) therefore transforms into:

$$y_{it} = \sum_{j=0}^{p} \Gamma_j [I_K \otimes \pi'(z_{it})] y_{i,t-j} + \epsilon_{it}$$

$$= \sum_{j=0}^{p} \Gamma_j x_{i,t-j} + \epsilon_{it} = \Gamma_j X_{i,t-1} + \epsilon_{it}, \tag{4.9}$$

where $X_{i,t-1} = (X'_{i,t-1}, X'_{i,t-2}, ..., X'_{i,t-p})'$ and $\Gamma = (\Gamma_1, \Gamma_2, ..., \Gamma_p)$. For the estimation purposes, the equation (4.9) can be rewritten as:

$$Y = \Gamma \cdot X + E, \tag{4.10}$$

where

$$Y = [y_{11}, y_{1T}, ...y_{N1}, ..., y_{NT}]_{(K \times NT)},$$

$$\Gamma = (\Gamma_1, \Gamma_2, ..., \Gamma_p)_{K \times pK\tau},$$

$$X = [X_{10}, X_{11}, ..., X_{1,T-1}, ..., X_{N0}, X_{N1}, X_{NT}]_{pK\tau \times NT},$$

and

$$E = [\epsilon_{11}, \epsilon_{12}, ..., \epsilon_{1T}, ..., \epsilon_{N1}, \epsilon_{N2}, ..., \epsilon_{NT}]_{K \times NT}.$$

Equation (4.10) in its vectorized form is:

$$y = (X' \otimes I_K) \cdot \gamma + \epsilon, \tag{4.11}$$

where $y = vec(Y)$, $\epsilon = vec(E)$ and $\gamma = vec(\Gamma)$. The OLS estimators are given by:

$$\hat{\Gamma} = YX'(XX')^{-1} \tag{4.12}$$

$$\hat{\gamma} = [(XX')^{-1}X \otimes I_K] \cdot y. \tag{4.13}$$

When allowing for cross-sectional heteroskedasticity, $\epsilon_{it} \overset{i.i.d.}{\sim} (0, \Sigma_{\epsilon,i})$, GLS estimation yields:

$$\hat{\gamma} = [(X \otimes I_K)S^{-1}(X' \otimes I_K)]^{-1}(X \otimes I_K)S^{-1}y, \tag{4.14}$$

where

$$S = \begin{bmatrix} I_T \otimes \Sigma_{\epsilon,1} & 0 & \dots & 0 \\ 0 & I_T \otimes \Sigma_{\epsilon,2} & \dots & 0 \\ \vdots & \ddots & \vdots & \\ 0 & 0 & \dots & I_T \otimes \Sigma_{\epsilon,N} \end{bmatrix}.$$

4.3.3 Impulse Responses

To investigate the properties of impulse responses in PCHVAR model, we rewrite the equation (4.4) in the following form:

$$\begin{bmatrix} y_{i,t} \\ y_{i,t-1} \\ \vdots \\ y_{i,t-p+1} \end{bmatrix} = \begin{bmatrix} A_1(z_{it}) & A_2(z_{it}) & \dots & A_{p-1}(z_{it}) & A_p(z_{it}) \\ I_K & 0 & \dots & 0 & 0 \\ 0 & I_K & & \vdots & 0 \\ \vdots & & \ddots & 0 & \vdots \\ 0 & 0 & \dots & I_K & 0 \end{bmatrix} \cdot \begin{bmatrix} y_{i,t-1} \\ y_{i,t-2} \\ \vdots \\ y_{i,t-p} \end{bmatrix} + \begin{bmatrix} \epsilon_{it} \\ 0 \\ \vdots \\ 0 \end{bmatrix},$$

which is equivalent to

$$Y_{it} = A(z_{it}) \cdot Y_{i,t-1} + E_{it}. \tag{4.15}$$

Recursive substitution into the equation (4.15) gives:

$$Y_{it} = A(z_{it}) \cdot A(z_{i,t-2}) \cdot Y_{i,t-2} + A(z_{it}) \cdot E_{i,t-1} + E_{it}$$

$$\vdots$$

$$= \big[\prod_{j=1}^{s} A(z_{i,t-j})\big] \cdot Y_{i,t-s} + \sum_{j=0}^{s} \big[\prod_{k=1}^{j} A(z_{i,t-k})\big] \cdot E_{i,t-j}. \tag{4.16}$$

Pre-multiplying the left side by $J_K = [I_K, 0, ..., 0]$ gives the following:

$$J_K \cdot Y_{it} = J_K \cdot \big[\prod_{j=1}^{s} A(z_{i,t-j})\big] \cdot Y_{i,t-s} + J_K \cdot \sum_{j=0}^{s} \big[\prod_{k=1}^{j} A(z_{i,t-k})\big] \cdot J_K' \cdot J_K \cdot E_{i,t-j}$$

$$y_{it} = \Upsilon_s \cdot Y_{i,t-s} + \sum_{j=0}^{s} \Phi_j(z_{i,t-j}^{(j)}) \cdot \epsilon_{i,t-j}, \tag{4.17}$$

where

$$\underline{z}_{i,t-1}^{j} = [z_{it,i,t-2}, ..., z_{i,t-j}]. \tag{4.18}$$

Therefore, the impulse responses are obtained from matrices:

$$\Phi_j(\underline{z}_{i,t-1}^{(j)}) = J_K \cdot [\prod_{k=1}^{j} A(z_{i,t-k})] \cdot J_K', \tag{4.19}$$

which implies that they are both history dependent and functions of the conditioning variables. Hence, we resort to the computation of generalized impulse responses.

4.4 Empirical Model of the CEE Economies

4.4.1 Data

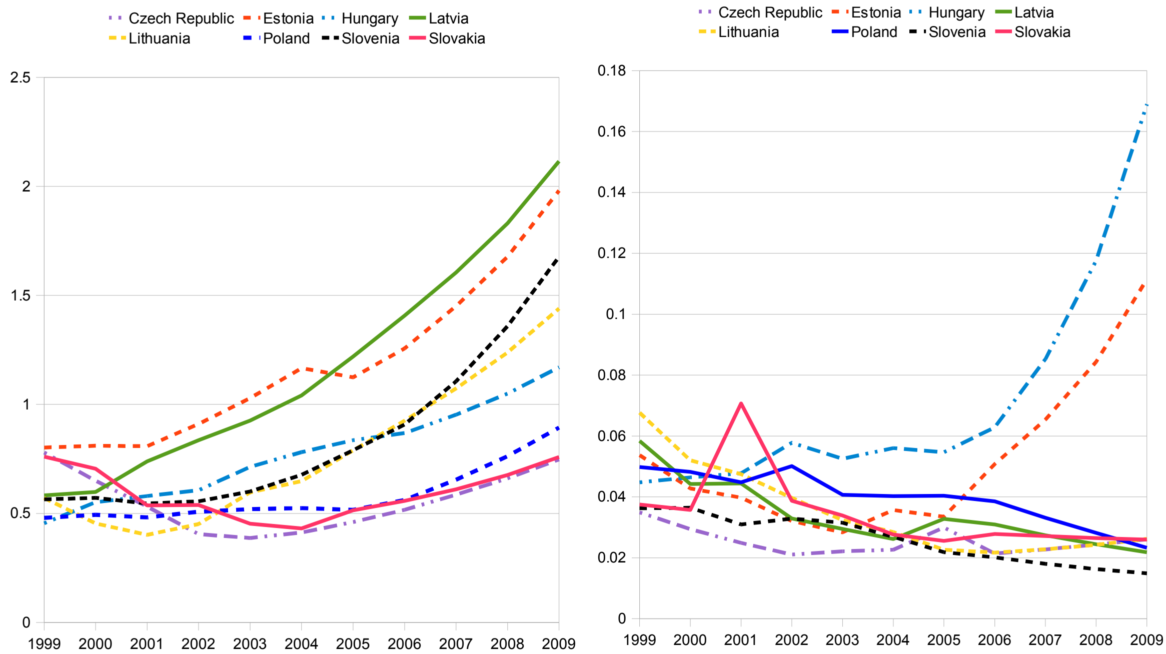
Our sample period spans between 1999:Q1 and 2009:Q4. Prior years witnessed high volatility and the adoption of new monetary policy frameworks. A danger of spurious correlations and misleading conclusions motivates our decision to restrict the sample. Post-2009 period is not considered, due to the issue of data availability.

Our specification is parsimonious and includes real GDP (*gdp*), price index (*pr*) and a three month money market rate (*ir*). All series were detrended using the HP filter ($\lambda = 1600$). We acknowledged the drawbacks of using generic values of λ in the Section 3.3.1. However, in an eight-country sample, with uncertain length of business cycles, standard value is the most sound solution. The series were retrieved from IMF International Financial Statistics (IFS) Database. Detailed description of the dataset is provided in the Table B.1.

In line with Georgiadis (2012b), we construct two indices aimed at reflecting the country's financial structure: credit dependence (*imp*) and banking sector competition (*comp*). Credit dependence is calculated by averaging series on bank credit relative to deposits and private credit relative to GDP. Assuming that more competitive banking sector features low interest margin and low costs, we construct the latter variable by averaging series on net interest margin

Figure 4.6: Financial structure indicators

(a) Credit dependence (b) Banking sector competition



and bank overhead costs to total assets.[5] The data on financial structure was obtained from World Bank Financial Structure Database (Beck *et al.* 2009). The series were converted to quarterly frequency using the quadratic-match procedure.

The newly constructed variables are plotted in the Figure 4.6. Individual country trends correspond to the stylized facts outlined in the Section 4.1. We observe an increasing credit dependence in all the countries since 1999. This ratio remains firmly below average in the Czech Republic, Poland and Slovakia throughout the period. On the other hand, it is rather high in Estonia and Latvia. The latter index indicates low competition at Hungarian and Estonian banking markets. This emanates from the high concentration in Estonian market and peculiarities of Hungarian banks, which charge high premium and simultaneously incur high overhead costs. The market appears somewhat less competitive in Poland, albeit no large-scale differences with the remaining countries are detected. The presence of trend in credit importance variable is evident in all the countries. Similarly, banking sector competition tightens across the sample, with the exception of Estonia and Hungary. We acknowl-

[5] For the estimation purposes, this series was subsequently inverted to ensure that higher values reflect more strenuous competition.

edge that the trend may plague our results by a higher degree of uncertainty; however, HP detrending of the four variables financial above is theoretically unappealing. Instead, we normalize the series to ensure cross-country comparability. Further, we attempt to validate the results using modified series.

The example of Estonia exposes the pitfalls of our analysis. High credit dependence ought to amplify the effects of monetary policy shocks, while low competition might dampen them. For technical reasons, a number of other relevant factors is not explicitly included, foreign lending being the most important one. Eventual impact of monetary policy may therefore be difficult to estimate.

4.4.2 Estimation Strategy

We recover the original shocks from reduced-form residuals using Cholesky decomposition. GDP is ordered first, followed by price index and interest rate. As outlined in the Chapter 2, this particular ordering implies that monetary authorities respond to contemporaneous changes in GDP and price level. However, no influence over their current values is exerted.

Our first venture is to estimate Linear VAR for each country in the sample. To test for the impact of monetary policy under distinct financial structures, the aforementioned financial variables are included as exogenous variables. Vectors of endogenous and exogenous variables respectively become

$$y_{1,t} = \begin{bmatrix} gdp_t & pr_t & ir_t \end{bmatrix}$$

and

$$y_{1,t} = \begin{bmatrix} imp_t & comp_t \end{bmatrix},$$

with the corresponding identification scheme

$$\begin{bmatrix} \epsilon_t^1 \\ \epsilon_t^2 \\ \epsilon_t^3 \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 \\ k_{21} & 1 & 0 \\ k_{31} & k_{32} & 1 \end{bmatrix} \begin{bmatrix} \mu_t^{gdp} \\ \mu_t^{pr} \\ \mu_t^{ir} \end{bmatrix}.$$

PCHVAR estimation is pursued employing both univariate and bivariate conditioning. In the former instance, impulse responses are computed over the grid of values of credit dependence and banking sector competition variables separately. Univariate conditioning hints about the role of one particular component of financial structure in monetary transmission. On the other hand, bivariate conditioning captures the interaction between two variables, rendering it indispensable for the purpose of more profound analysis. Impulse responses are computed by fixing one conditioning variable, while computing the responses over a grid of values of the remaining one.

4.4.3 Results

We first estimate the Linear VAR for each country in the sample. The estimation is carried out with the detrended variables and one lag. Bearing in mind the degrees of freedom restrictions, this particular lag length is selected to ensure cross-country comparability. AR characteristic polynomial roots indicate that all the country-specific models are stable.[6]

Single-country estimation output is given in the Figure B.1. Inspection of the impulse responses gives some preliminary insights about the role of financial structure in monetary transmission. Drawing uniform conclusions about an entire sample is tedious, however. GDP responds negatively to monetary tightening in most of the countries. An exception is Slovakia, where this reaction is positive. GDP tends to bottom out between the fifth and the tenth quarter. The maximum response of GDP is comparable between the five countries, Hungary, Poland and Slovakia being an exception. In the latter subsample, the response is small or insignificant. It is worth noting that in the cases of Estonia, Latvia and Lithuania, the impulse responses do not die out completely even after 30 quarters. However, the confidence intervals are fairly wide. Resembling impulse responses of prices are evident, again with the exception of Poland, Hungary and Slovakia. Prices in Slovakia fall in the wake of monetary contraction; however, they bottom out almost immediately to return back to zero fairly quickly. No considerable differences in terms of interest rate responses can be detected.

Next, we test for the impact of financial structure variables on monetary transmission within the framework of PCHVAR. The model is estimated on one lag, merely to conserve the degrees of freedom and ensure comparability with Linear VAR. Stability is not explicitly checked for; however, in our baseline model, impulse responses die out after a certain period. This observation represents an indirect proof of stability.

[6] These results are available upon request.

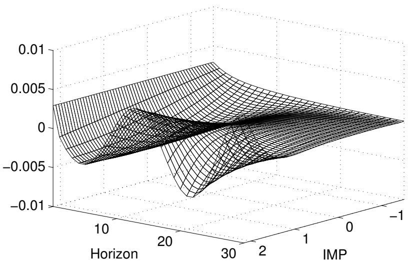
Impulse responses obtained from PCHVAR are by no means indicators of the overall effectiveness of monetary policy. Instead, they merely demonstrate what the cross-country differences in monetary transmission would be if they diverged in one particular aspect of financial structure, everything else being the same. Our estimation begins with the univariate conditioning. Impulse responses are generalized, which provides additional information about the interaction between the series, without imposing severe restrictions on the coefficients. Impulse responses from Figure 4.7 indicates that the reaction of GDP, prices and interest rate strengthens at higher values of credit dependence variable. Further, we can infer that at the very low values of this variable, monetary transmission breaks down. On the other hand, impulse responses over the grid of banking sector competition in the Figure 4.8 are counterintuitive: more strenuous competition appears to dampen the effects of monetary policy on both GDP and prices. Its potency increases with the decline in competitive pressures.

We discus the results in the context of interest rate and credit channel. Higher importance of credit tends to amplify the effects of monetary shocks. Our results are in line with this observation. Credit dependence is the highest in the Baltics, followed by Slovenia and Hungary. Indeed, Köhler *et al.* (2005) and Beņkovskis (2008) argue that this particular feature fosters the bank lending channel in the three Baltic states. Similar claim has been made for Slovenia (Ahtik 2012) and Hungary (Horváth *et al.* 2006). On the other hand, Łyziak *et al.* (2008) argue that high liquidity in the Polish banking sector renders this channel dysfunctional. The evidence on the Czech Republic is less clear-cut. Credit dependence index is among the lowest in the sample. Counterintuitively, Fidrmuc *et al.* (2008) and Matousek & Sarantis (2009) find that both balance sheet and bank lending channel are operative. However, their practical relevance is not evaluated. Contrary to these, Niedermayer (2008) notes that the interest rates on mortgage loans tend to be fixed, which may render less responsiveness to policy rate shifts in the banking sector. It is worth highlighting that the cross-country comparisons on the extent of credit channel functionality are inherently difficult to draw. For more intuition, we turn to analyze the other results.

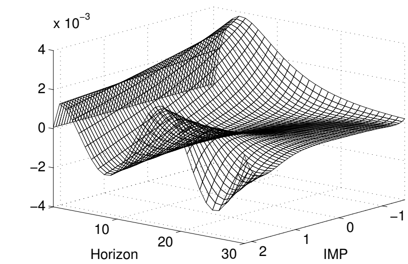
Mainstream theoretical studies suggest that higher degree of competition fosters monetary transmission. Our results do not corroborate with observation. However, we do not diverge considerably from the other studies on CEE region. Competition index indicates that the Czech Republic, Slovenia and

Figure 4.7: Impulse responses from PCHVAR conditioned on credit dependence

(a) GDP



(b) Prices



(c) Interest rate

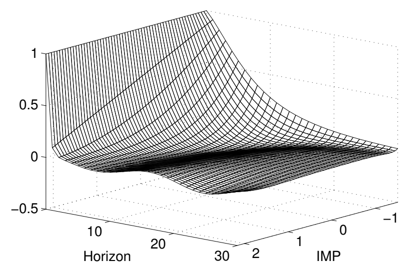
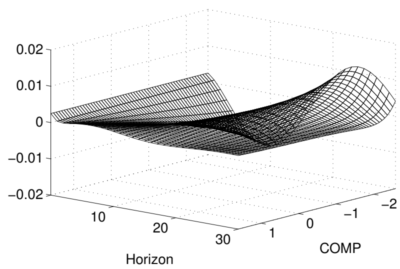
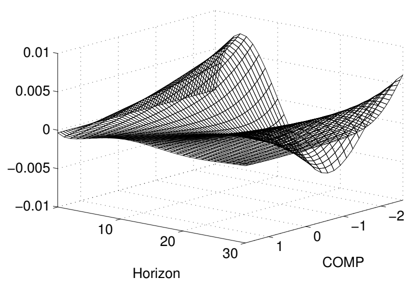


Figure 4.8: Impulse responses from PCHVAR conditioned on banking competition

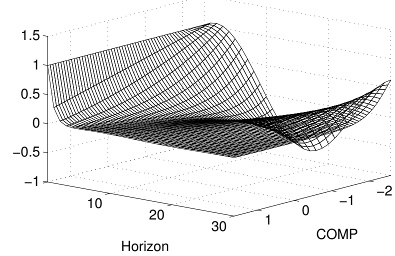
(a) GDP



(b) Prices



(c) Interest rate



Slovakia have the most competitive market, followed by Lithuania, Latvia and Poland, while Estonia and Hungary are at the bottom. Some positive correlation between the competition and the pass-through notwithstanding, Kot (2004) finds that Poland and Hungary exhibit higher degree of competition and weaker pass-through than the euro zone average. Hence, this mechanism may not be entirely operative. Next, Égert *et al.* (2007) and Égert & MacDonald (2008) provide evidence of a weak pass-through in the Czech Republic and Slovakia, which is contrary to that in the Baltics. The results for Hungary are less clear-cut. Horváth *et al.* (2004) argue that the interest rate channel in Hungary is functional, in spite of downward rigidity of the corporate rates. Égert & MacDonald (2008) indicate that the pass-through to household lending rates is weak. Thus, the ultimate repercussion of competition on monetary policy may be difficult to evaluate. The results for Slovenia are ambiguous across the studies. Further, Baltic states and Hungary have relatively higher share of corporate loans, which are more sensitive to policy shifts.

Policy advice on one particular aspect of financial structure may be of a limited relevance, due to the fact that different policy measures act in conjunction. Specifically, different components of financial structure may offset each other, which has implications on previously discussed results. Consider the cases of Estonia and Latvia. In 2008, credit-to-GDP ratio reached 91.9% and 89.6% respectively, which is an indicator of high credit dependence. This, combined with the high share of corporate loans and negligible government presence in the banking sector, ought to increase the leverage by monetary authorities. However, foreign subsidiaries, which own nearly 100% of the domestic banks, may circumvent the effects of monetary contraction by borrowing from the mother companies. All these factors, combined with a high concentration and high amount of foreign lending, render the eventual impact of monetary policy unclear. On the other side of the spectrum, the Czech Republic is the least credit dependent country, which impedes the functioning of monetary policy. However, negligible foreign lending and a relatively low share of corporate loans ought to bolster its effectiveness.

The role of financial structure in monetary transmission therefore needs to be evaluated employing a range of indicators. In the context of PCHVAR, such an estimation may be technically difficult to carry out. Hence, we resort to bivariate conditioning, as described in the Section 4.3.2. In comparison with the previous scenario, the results presented in Figure B.3 and Figure B.2 provide a different picture. Impulse responses computed over the grid of banking sector competition variable, while credit dependence is fixed at mean, indicates negligible divergences in monetary transmission at different values. The picture is nearly the same when impulse responses are computed over the grid of credit dependence variable. In both instances, GDP bottoms out around the eighth and prices in the tenth quarter. The responses die out between the fifteenth and the twentieth quarter. As discussed in the Chapter 3, price puzzle probably emerges due to the fact that no variables capturing external developments have been included.

The above presented results yield an important conclusion: when interacted with each other, different values of credit dependence or banking sector competition do not alter the effectiveness of monetary policy. This confirms our previous suspicion that various components of financial structure have a mutually offsetting impact, leveling off the cross-country variability. Assenmacher-Wesche & Gerlach (2008) put forward the same conclusion in the context of OECD economies: Panel VAR does not implicate the financial structure as a significant determinant of monetary transmission, which is precisely due to the counterbalancing effects described above. We acknowledge that our findings are contestable. Specifically, the indicators of credit dependence and competition are constructed as a combination of two different time series. As Elbourne & De Haan (2004) argue, this procedure is of subjective and *ad hoc* nature, which may generate misleading conclusions. However, conclusions based on individual components of financial structure are of a limited informative value, as the ultimate impact of policy needs to be evaluated.

4.5 Sensitivity Analysis

We perform the sensitivity check with annual time series. In the first part, financial variables were interpolated from annual to quarterly frequency. We acknowledged potential pitfalls of our strategy above. However, this decision was based on the fact that the annual sample is too short, which may substantially restrict the degrees of freedom and render imprecise impulse responses.

The results of univariate conditioning are presented in Figure B.4 and Figure B.5. Monetary transmission appears more potent under the conditions of higher credit dependence and weaker competition. This finding broadly confirms our initial results. It should be noted that the impulse responses at very low levels of each variable do not die out. This is indicative about sparsely populated grid space in this region. Hence, these particular responses may be imprecise. Next, the results also hold for bivariate conditioning: the responses are broadly uniform across the grid of values of conditioning variables. However, a slight increase in the strength of response can be observed at very low values of competition. Although this finding may look counterintuitive, we have demonstrated that different segments of financial structure interact with each other, which may result in unexpected responses. Moreover, the ultimate impact of monetary policy should be judged using a variety of financial structure indicators. In this instance, the unobserved component (high share of FX loans, predominant foreign ownership) works to diminish the effects of monetary policy in the economies characterized by strong competition.

Chapter 5

Conclusion

In this thesis, we studied monetary transmission asymmetries in CEE region. The first part is focused on the Czech Republic. Using the Logistic Smooth Transition Vector Autoregression (LSTVAR) model for the 1998:M1-2012:M3 time span, we investigate the effectiveness of monetary policy at different stages of the credit cycle. In the second part, the role of financial structure in monetary transmission is examined. This analysis is conducted within the framework of Panel Conditionally Homogenous Vector Autoregression (PCHVAR) model, on a sample of eight CEE economies – the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Slovenia and Slovakia for the period spanning between 1999:Q1 and 2009:Q4.

The main conclusions are as follows. First, credit cycle determines the effectiveness of monetary policy in the Czech Republic. Credit growth above the threshold is found to impair the monetary authorities' leverage. To rephrase, in the environment of abundant credit liquidity, agents have sufficient funds to counter the negative effects of monetary contraction. On the other hand, deteriorating financial conditions amplify the effects of monetary tightening. No asymmetries in relative effects of contractionary and expansionary shocks have been documented. This may emanate from the fact that our sample encompasses a single monetary policy regime, i.e. post-1998 period. Our results indicate that the broad credit channel in the Czech Republic is operative.

Second, certain aspects of financial structure may induce monetary transmission asymmetries. Higher credit dependence is found to foster the interest rate pass-through in CEE region. On the other hand, monetary policy appears less potent if the banking sector competition is strong. This is clearly counterintuitive, as more strenuous competition ought to result in more susceptibility to the outside pressures. However, individual country experiences are more nuanced. Hungary and Estonia have the least competitive banking market, but also a high share of corporate loans, which are more sensitive to monetary policy shifts. On the other side of the spectrum, the Czech Republic has a competitive market with a lower share of corporate loans.

Third, the ultimate impact of financial structure on monetary transmission is indeterminable. When the credit dependence and the banking competition are interacted with each other, the evidence of asymmetries largely disappears. Again, individual experiences are revealing. Baltic states are highly dependent on credit, which makes them more susceptible to monetary actions. On the other hand, their banking markets are very concentrated. More importantly, these countries engaged in excessive FX borrowing, over which the domestic monetary authorities have virtually no influence. Therefore, the individual components of financial structure may offset each other's effects.

The ultimate role of financial structure in monetary transmission is difficult to asses for three reasons. First, capturing its segments in a comprehensive manner is tedious. Second, few econometric techniques have been developed for this type of research. Further, mutual effects of monetary and prudential policies need to be examined. Considering its incontestable importance in policy-making process, the research on financial structure needs to address these issues.

Bibliography

Ahtik, M. (2012): "Bank Lending Channel in Slovenia: Panel Data Analysis." *Prague Economic Papers* **(1)**.

Arnoštová, K. & J. Hurník (2005): "The Monetary Transmission Mechanism in the Czech Republic (evidence from VAR analysis)." *Working Paper Series 4*, Czech National Bank.

Assenmacher-Wesche, K. & S. Gerlach (2008): "Ensuring Financial Stability: Financial Structure and the Impact of Monetary Policy on Asset Prices." *Working Paper 361*, Institute for Empirical Research in Economics, University of Zurich.

Atanasova, C. (2003): "Credit Market Imperfections and Business Cycle Dynamics: A Nonlinear Approach." *Studies in Nonlinear Dynamics and Econometrics* **7(4)**: p. 5.

Azariadis, C. & B. Smith (1998): "Financial Intermediation and Regime Switching in Business Cycles." *The American Economic Review* **88(3)**: pp. 516–536.

Babetskaia-Kukharchuk, O. (2007): "Transmission of Exchange Rate Shocks into Domestic Inflation: The Case of the Czech Republic." *Working Paper Series 12*, Czech National Bank.

Balke, N. (2001): "Credit and Economic Activity: Credit Regimes and Nonlinear Propagation of Shocks." *Review of Economics and Statistics* **82(2)**: pp. 344–349.

Ball, L. & G. Mankiw (1994): "Asymmetric Price Adjustment and Economic Fluctuations." *The Economic Journal* **(104)**: pp. 247–261.

Bathaluddin, M. B., N. M. Adhi, & A. Wahyu (2012): "The Impact of Excess Liquidity on Monetary Policy." *Bank of Indonesia Economic Journal* **14(3)**.

Beck, T., A. Demirgüç-Kunt, & R. Levine (2009): "Financial Institutions and Markets Across Countries and over Time: Data and Analysis." *World Bank Policy Research Working Paper 4943*, World Bank.

Beņkovskis, K. (2008): "Is there a bank lending channel of monetary policy in Latvia? Evidence from bank level data." *Technical Report 1*, Latvian National Bank.

Bernanke, B. & A. Blinder (1988): "Credit, Money and Aggregate Demand." *Americal Economic Reivew* **78(2)**: pp. 435–439.

Bernanke, B. & M. Gertler (1989): "Agency Costs, Net Worth and Business Fluctuations." *The American Economic Review* **79(2)**: pp. 14–31.

Bernanke, B. & M. Gertler (1995): "Inside the Black Box: The Credit Channel of Monetary Policy Transmission." *Journal of Economic Perspectives* **9(4)**: pp. 27–48.

Bernanke, B., M. Gertler, & S. Gilchrist (1996): "The Financial Accelerator and the Flight to Quality." *The Revew of Economics and Statistics* **78(1)**: pp. 1–15.

Bernanke, B., M. Gertler, & S. Gilchrist (1998): "The Financial Accelerator in a Quantitative Business Cycle Framework." *NBER Working Paper 6455*, National Bureau of Economic Research.

Blinder, A. S. (1987): "Credit Rationing and Effective Supply Failures." *The Economic Journal* **(97)**: pp. 327–352.

Bordon, A. R. & A. Weber (2010): "The Transmission Mechanism in Armenia: New Evidence from a Regime Switching VAR Analysis." *IMF Working Paper 270*, International Monetary Fund.

Borys, M. M., M. Franta, & R. Horváth (2009): "The effects of monetary policy in the Czech Republic: an empirical study." *Empirica* **36(4)**: pp. 419–443.

CALZA, A. & J. SOUSA (2005): "Output and Inflation Responses to Credit Shocks: Are there Threshold Effects in the Euro-area?" *Working paper 481*, European Central Bank.

CANOVA, F. (2006): "You can use VARs for structural analyses. A comment to VARs and the Great Moderation." *Technical report*, ICREA, Universitat Pompeu Fabra, CREI and CEPR.

CANOVA, F. (2007): *Methods for Applied Macroeconomic Research.* Princeton University Press.

ÇATIK, A. N. & C. MARTIN (2012): "Macroeconomic transitions and the transmission mechanism: Evidence from Turkey." Forthcoming.

CECCHETTI, S. G. (1999): "Legal Structure, Financial Structure, and the Monetary Policy Transmission Mechanism." *FRBNY Economic Policy Review*, Federal Reserve Bank of New York.

CHARI, V., P. J. KEHOE, & E. MCGRATTAN (2008): "Are Structural VARs with Long-run Restrictions Useful in Developing Business Cycle Theory?" *NBER Working Paper 14430*, National Bureau of Economic Research.

CHRISTIANO, L., M. EICHENBAUM, & C. L. EVANS (1999): "Monetary policy shocks: What have we learned and to what end?" In J. TAYLOR & M. WOODFORD (editors), "Handbook of Macroeconomics," volume 1, chapter 2, pp. 65–148. Elsevier.

COGLEY, T. & T. SARGENT (2001): "Evolving Post-World War II U.S. Inflation Dynamics." In "NBER Macoreconomics Annual," volume 16, pp. 331–338. National Bureau of Economic Research.

COGLEY, T. & T. SARGENT (2005): "Drifts and Volatilities: Monetary Policies and Outcomes in the Post WWII U.S." *Review of Economic Dynamics* **8(2)**: pp. 262–302.

COOPER, R. & D. CORBAE (2001): "Financial Collapse and Active Monetary Policy: A Lesson from the Great Depression." *Staff Report 289*, Federal Reserve Bank of Minneapolis.

COVER, J. (1988): "A Keynesian Macroeconomic Model with New-Classical Econometric Properties." *Southern Economic Journal* **(54)**: pp. 831–839.

Cushman, D. & T. Zha (1997): "Identifying monetary policy in a small open economy under flexible exchange rates." *Journal of Monetary Economics* **39**: pp. 433–448.

Darvas, Z. (2009): "Monetary Transmission in three Central European Economies: Evidence from Time-Varying Coefficient Vector Autoregressions." *DNB Working Papers 208*, Netherlands Central Bank.

De Bondt, G. (2005): "Interest Rate Pass-Through: Empirical Results for the Euro Area." *German Economic Review* **6(1)**: pp. 37–78.

De Long, J. B. & L. H. Summers (1988): "How does Macroeconomic Policy Affect Output?" In W. C. Brainard & G. L. Perry (editors), "Brookings papers on economic activity," pp. 433–494.

Demel, W. & J. Sikimic (2008): "CEE Banking Sector Report." *Technical report*, Raiffeisen Research.

EBRD (2009): "Transition in Crisis." *Transition Report 09*, European Bank for Reconstruction and Development.

Economic and Financial Affairs Directorate General (2010): "Convergence Report."

Égert, B., P. Backé, & T. Zumer (2006): "Credit Growth in Central and Eastern Europe: New (Over)Shooting Stars?" *Working Paper Series 687*, European Central Bank.

Égert, B., J. Crespo-Cuaresma, & T. Reininger (2007): "Interest Rate Pass-Through in Central and Eastern Europe: Reborn from Ashes Merely to Pass Away?" *Journal of Policy Modeling* **29(2)**: pp. 209–225.

Égert, B. & R. MacDonald (2008): "Monetary Transmission Mechanism in Central and Eastern Europe: Surveying the Surveyable." *OECD Economics Department Working Papers 654*, OECD Publishing.

Ehrmann, M., L. Gambacorta, J. Martínez Pagés, P. Sevestre, & A. Worms (2001): "Financial Systems and the Role of Banks in Monetary Policy Transmission in the Euro Area." *Working Paper 105*, European Central Bank.

Elbourne, A. & J. De Haan (2004): "Asymmetric Monetary Transmission in EMU: The Robustness of VAR Conclusions and Cecchetti's Legal Family Theory." *CESifo Working Paper Series 1327*, CESifo Group Munich.

Elbourne, A. & J. De Haan (2006): "Financial structure and monetary policy transmission in transition countries." *Journal of Comparative Economics* **34(1)**: pp. 1–23.

Enders, W. (2004): *Applied Econometric Time Series*. Wiley, second edition.

Farinha, L. & C. Robalo Marques (2001): "The Bank Lending Channel of Monetary Policy: Identification and Estimation Using Portugese Micro Bank Data." *Working Paper 102*, European Central Bank.

Fernández-Villaverde, J., J. F. Rubio-Ramírez, & T. Sargent (2005): "A, B, C's (and D's for Understanding VARs)." *Working paper 9*, Federal Reserve Bank of Atlanta.

Fidrmuc, J., R. Horváth, & E. Horváthová (2008): "Corporate Interest Rates and the Financial Accelerator in the Czech Republic." *Discussion Papers in Economics 7191*, University of Munich, Department of Economics.

Frait, J., A. Geršl, & J. Seidler (2011): "Credit Growth and Financial Stability in the Czech Republic." *Policy Research Working Paper 5771*, The World Bank.

Franta, M., R. Horváth, & M. Rusnák (2012): "Evaluating Changes in the Monetary Transmission Mechanism in the Czech Republic." *IES Working Paper 11*, Charles University in Prague.

Fuchun, L. (2010): "Financial Stress, Monetary Policy, and Economic Activity." *Bank of Canada Review*, Bank of Canada.

Galbraith, J. (1996): "Credit Rationing and Threshold Effects in the Relation Between Money and Output." *Journal of Applied Econometrics* **11(4)**: pp. 419–429.

Galvao, A. B. & M. Marcellino (2011): "An endogenous threshold VAR model of the monetary transmission mechanism." Under revision for invited resubmission at the Studies for Nonlinear Dynamics and Econometrics.

Garcia, R. & H. Schaller (2002): "Are the Effects of Monetary Policy Asymmetric?" *Economic Inquiry* **40(1)**: pp. 102–119.

Georgiadis, G. (2012a): "The Panel Conditionally Homogenous Vector Autoregressive Model." *MPRA Paper 37755*, University Librarty of Munich.

Georgiadis, G. (2012b): "Towards an Explanation of Cross-Country Asymmetries in Monetary Transmission." *Discussion Paper 07/2012*, Deutsche Bundesbank.

Gertler, M. & S. Gilchrist (1994): "Monetary Policy, Business Cycles and the Behavior of Small Manufacturing Firms." *The Quarterly Journal of Economics* **109(2)**: pp. 309–340.

Gilchrist, S., V. Yankov, & E. Zakrajsek (2009): "Evidence from the Composition of External Finance." *NBER Working Paper 14863*, National Bureau of Economic Research.

Giordani, P. (2001): "An Alternative Explanation of the Price Puzzle." *Sveriges Riksbank Working Paper Series 125*, Sveriges Riksbank.

Goodhart, C. & B. Hofmann (2008): "House Prices, Money, Credit and the Macroeconomy." *Working Paper 888*, European Central Bank.

Gropp, R., C. Kok Sørensen, & J. Lichtenberger (2007): "The Dynamics of Bank Spreads and Financial Structure." *Technical Report 714*, European Central Bank.

Harvey, A. & A. Jaeger (1993): "Detrending, Stylized Facts and the Business Cycle." *Journal of Applied Econometrics* **8(3)**: pp. 231–247.

Havránek, T., J. Matějů, & R. Horváth (2010): "Monetary transmission and the financial sector in the Czech Republic." *Working Paper 06*, Czech National Bank.

Hloušek, M. (2006): "Czech Business Cycle Stylized Facts." *Working Paper 10*, Research Centre for Competitiveness of Czech Economy.

Holmstrom, B. & J. Tirole (1997): "Financial Intermediation, Loanable Funds and the Real Sector." *The Quarterly Journal of Economics* **112(3)**: pp. 663–691.

Horváth, C., J. Krekó, & A. Naszódi (2004): "Interest rate pass-through in Hungary." *MNB Working Papers 2004/8*, Hungarian National Bank.

Horváth, C., J. Krekó, & A. Naszódi (2006): "Is there a bank lending channel in Hungary? Evidence from bank panel data." *MNB Working Papers 7*, Hungarian National Bank.

Horváth, R. Rusnák, M. (2008a): "How Important Are Foreign Shocks in Small Open Economy? The Case of Slovakia." *William Davidson Institute Working Paper 933*, The William Davidson Institute.

Horváth, R. (2008b): "Asymmetric Monetary Policy in the Czech Republic?" *Czech Journal of Economics and Finance* **58(9-10)**.

Huang, Z. (2003): "Evidence of a bank lending channel in the UK." *Journal of Banking & Finance* **27(3)**: pp. 491–510.

Jääskelä, J. (2007): "More Potent Monetary Policy? Insights from a Threshold Model." *Research Discussion Paper 07*, Reserve Bank of Australia.

Jarociński, M. (2008): "Responses to Monetary Policy Shocks in the East and the West of Europe: A Comparison." *Working Paper 970*, European Central Bank.

Jiménez-Rodriguez, R., A. Morales-Zumaquero, & B. Égert (2010): "The VARying Effect of Foreign Shocks in Central and Eastern Europe." *Working Paper 3080*, CESIFO.

Juks, R. (2004): "The Importance of the Bank-Lending Channel in Estonia: Evidence from Micro-Economic Data." *Working Papers of Eesti Pank 6*, Estonian National Bank.

Kandil, M. (1995): "Asymmetric Nominal Flexibility and Economic Fluctuations." *Southern Economic Journal* **(61)**: pp. 674–695.

Karamé, F. & A. Olmedo (2002): "The Asymmetric Effects of Monetary Policy Shocks: A Nonlinear Structural VAR Approach." Preliminary version.

Kashyap, A. & J. Stein (1994): "The Impact of Monetary Policy on Bank Balance Sheets." *NBER Working Paper 4821*, National Bureau of Economic Research.

Kashyap, A. & J. Stein (2000): "What Do a Million Observations on Banks Say about the Transmission of Monetary Policy?" *Americal Economic Reivew* **(3)**: pp. 407–428.

Kashyap, A., J. Stein, & D. Wilcox (1993): "Monetary Policy and Credit Conditions: Evidence from the Composition of External Finance." *The American Economic Review* **83(1)**: pp. 78–98.

Kishan, R. & T. Opiela (2000): "Bank Size, Bank Capital and Bank Lending Channel." *Journal of Money, Credit and Banking* **32(1)**: pp. 121–141.

Kiyotaki, N. & J. Moore (1997): "Credit Cycles." *The Journal of Political Economy* **105(2)**: pp. 211–248.

Köhler, M., J. Hommel, & M. Grote (2005): "The Role of Banks in the Transmission of Monetary Policy in the Baltics." *Discussion Paper 6*, Zentrum für Europäische Wirtschaftsforschung.

Koop, G. (2003): *Bayesian econometrics*. Wiley.

Koop, G., H. M. Pesaran, & S. M. Potter (1996): "Impulse response analysis in nonlinear multivariate models." *Journal of Econometrics* **(74)**: pp. 119–147.

Kot, A. (2004): "Does Banking Sector Competitiveness Affect Interest Rate Pass-Through?" Paper presented at the 3rd Macroeconomic Policy Research Workshop on October 29-30, 2004. National Bank of Hungary.

Kowal, P. (2005): "Matlab implementation of commonly used filters." *Econ-WPA Computer Programs* **(0507001)**.

Kwapil, C. & J. Scharler (2006): "Limited Pass-Through from Policy to Retail Interest Rates: Empirical Evidence and Macroeconomic Implications." In "Monetary Policy & the Economy," volume Q4
06, pp. 26–36.

Lo, M. & J. Piger (2003): "Is the Response of Output to Monetary Policy Asymmetric? Evidence from a Regime-Switching Coefficients Model." *Working Paper 2001-022D*, Federal Reserve Bank of St. Louis.

Łyziak, T., J. Przystupa, & E. Wróbel (2008): *Monetary Policy Transmission in Poland: A Study of the Importance of Interest Rate and Credit*

Channels, chapter 1, pp. 7–69. Chapters in SUERF Studies. SUERF - The European Money and Finance Forum.

Maćkowiak, B. (2005): "How much of the Macroeconomic Variation in Eastern Europe is Attributable to External Shocks?" *SFB 649 Discussion Paper 061*, Humboldt-Universität zu Berlin.

Mandler, M. (2010): "Macroeconomic dynamics and inflation regimes in the U.S. Results from threshold vector autoregressions." *MPRA Paper 21887*, University Library of Munich.

Matousek, R. & N. Sarantis (2009): "The bank lending channel and monetary transmission in Central and Eastern European countries." *Journal of Comparative Economics* **37(2)**: pp. 321–334.

McCallum, J. (1991): "Credit Rationing and the Monetary Transmission Mechanism." *The American Economic Review* **81(4)**.

MNB (2009): "Report on Financial Stability." *Technical report*, Hungarian National Bank.

Mojon, B. (2000): "Financial Structure and the Interest Rate Channel of ECB Monetary Policy." *Working Paper 40*, European Central Bank.

Mojon, B. & G. Peersman (2001): "A VAR Description of the Effects of Monetary Policy in the Individual Countries of the Euro area." *Working Paper 092*, European Central Bank.

Moody, A. & M. Taylor (2003): "The High Yield Spread as a Predictor of Real Economic Activity: Evidence of a Financial Accelerator for the United States." *IMF Working Papers 50(3)*, International Monetary Fund.

Niedermayer, L. (2008): "Notes on the monetary transmission mechanism in the Czech economy." In "Transmission mechanisms for monetary policy in emerging market economies," BIS Papers No. 35, pp. 213–219. Bank for International Settlements.

Peersman, G. & F. Smets (2001): "Are the Effects of Monetary Policy in the Euro-area Greater in Recessions Than in Booms?" *Working Paper 52*, European Central Bank.

Poměnková, J. & R. Maršálek (2011): "Time and frequency domain in the business cycle structure." *MENDELU Working Papers in Business and Economics 7*, Mendel University in Brno.

Primiceri, G. (2005): "Time Varying Structural Vector Autoregression and Monetary Policy." *Review of Economic Studies* **(72)**: pp. 821–852.

Pruteanu, A. (2004): "The Role of Banks in the Czech Monetary Transmission Mechanism." *Working Paper Series 3*, Czech National Bank.

Ravn, M. & M. Sola (1999): "Asymmetric Effects of Monetary Policy in the US: Positive versus Negative or Big versus Small?" *Working paper*, Universitat Pompeu Fabra.

Rousseas (1985): "A Markup Theory of Bank Loan Rates." *Journal of Post Keynesian Economics* **8(1)**: pp. 135–144.

Rusnák, M., T. Havránek, & R. Horváth (2011): "How to Solve the Price Puzzle? A Meta-Analysis." *IES Working Paper 24*, Charles University in Prague.

Sá, F., P. Towbin, & Wieladek (2011): "Low Interest Rates and Housing Booms: the Role of Capital Inflows, Monetary Policy and Financial Innovation." *Bank of England Working Papers 411*, Bank of England.

Saxegaard, M. (2006): "Excess Liquidity and Effectiveness of Monetary Policy: Evidence from Sub-Saharan Africa." *IMF Working Paper 115*, International Monetary Fund.

Shen, C. & T. Chiang (1999): "Retrieving the Vanishing Liquidity Effect - A Threshold Vector Autoregressive Model." *Journal of Economics and Business* **(51)**: pp. 259–277.

Sims, C. & T. Zha (1998): "Does Monetary Policy Generate Recessions?" *Working Paper 12*, Federal Reserve Bank of Atlanta.

Sims, C. A. (1980): "Macroeconomics and Reality." *Econometrica* **48(1)**: pp. 1–48.

Smith, R. P. & A. M. Fuertes (2010): "Panel Time-Series."

Stiglitz, J. & A. Weiss (1981): "Credit Rationing in Markets with Imperfect Information." *The American Economic Review* **71(3)**: pp. 393–410.

STOCK, J. H. & M. W. WATSON (2001): "Vector Autoregressions." *Journal of Economic Perspectives* **15(4)**: pp. 101–115.

THOMA, M. A. (1994): "Subsample Instability and Asymmetries in Money-Income Causality." *Journal of Econometrics* **(64)**: pp. 279–306.

TSIDDON, D. (1993): "The (Mis)Behavior of the Aggregate Price Level." *Review of Economic Studies* **(60)**: pp. 889–902.

VAN LEUVENSTEIJN, M., C. KOK SØRENSEN, J. A. BIKKER, & A. VAN RIXTEL (2008): "Impact of Bank Competition on the Interest Rate Pass-Through in the Euro Area." *Working Paper Series 885*, European Central Bank.

VAŠÍČEK, B. (2011): "Is Monetary Policy in the New EU Member States Asymmetric." *Working Paper Series 5*, Czech National Bank.

WEISE, C. L. (1999): "The asymmetric effects of monetary policy: A nonlinear vector autoregression approach." *Journal of Money, Credit and Banking* **31(1)**: pp. 85–108.

WETH, M. (2002): "The pass-through from market interest rates to bank lending rates in Germany." *Discussion Paper 11*, Economic Research Centre of the Deutsche Bundesbank.

WRÓBEL, E. & M. PAWŁOWSKA (2002): "Monetary transmission in Poland: some evidence on interest rate and credit channels." *Technical Report 24*, National Bank of Poland, Economic Institute.

Appendix A

LSTVAR Data and Estimation Output

The CD enclosed to this thesis contains empirical data and Matlab codes, along with the instructions for their use.

- Folder 1: Matlab codes + instructions
- Folder 2: Empirical data

Table A.1: Dataset

Variable	Unit	Source
GDP, s.a.	EUR mil	IMF IFS
industrial production, s.a.	2005=100	IMF IFS
consumer price index net	2005=100	IMF IFS
3 month money market rate	p.a.	IMF IFS
total credits	CZK mil	World Bank
PX index	end of period	Prague Stock Exchange
EUR/CZK nominal exchange rate	monthly average	Eurostat
Euro zone (EA-12) GDP, s.a.	EUR mil.	Eurostat
Euro zone (EA-12) HICP	2005=100	Eurostat
Euro zone (EA-12) money market rate	EUR mil.	Eurostat

Figure A.1: Time series used in estimation, HP detrended

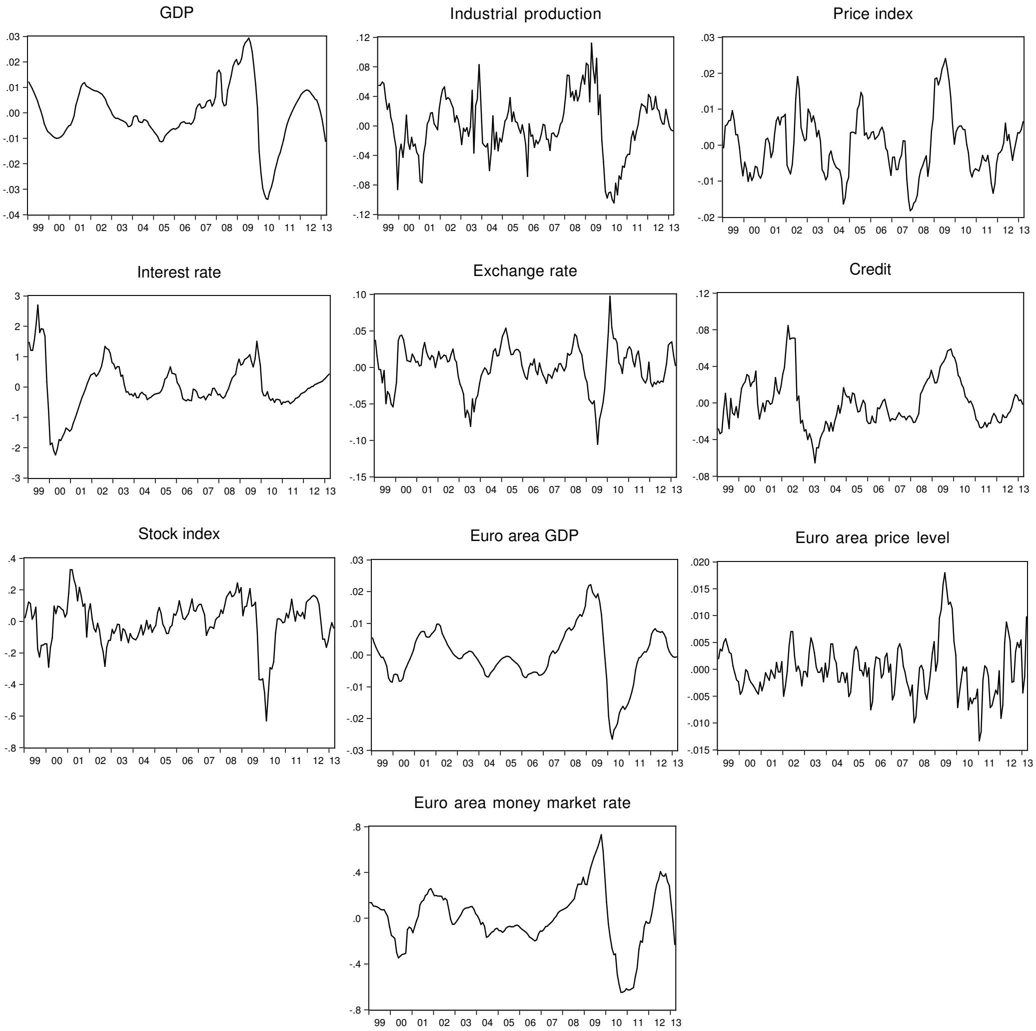


Table A.2: Lag length selection criteria for final specification

Lag	LogL	LR	FPE	AIC	SC	HQ
0	1576.444	NA	2.18e-16	-19.03568	-18.92274	-18.98984
1	2376.056	1531.378	2.08e-20	-28.29158	-27.50098*	-27.97065*
2	2426.278	92.53096*	1.75e-20*	-28.46398*	-26.99571	-27.86796
3	2452.884	47.08503	1.97e-20	-28.35011	-26.20419	-27.47901
4	2473.574	35.10934	2.40e-20	-28.16453	-25.34094	-27.01834
5	2504.183	49.71749	2.59e-20	-28.09919	-24.59795	-26.67792
6	2528.025	36.99094	3.06e-20	-27.95182	-23.77291	-26.25546

Figure A.2: AR roots of characteristic polynomial for final specification

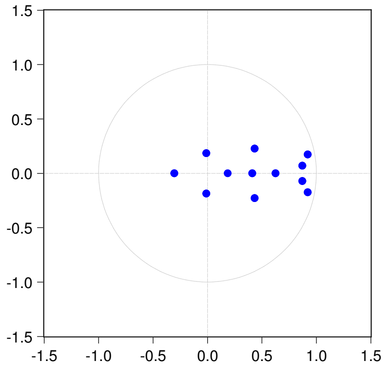


Table A.3: Granger Causality/Block Exogeneity Test, p-values

Excluded (x)	**Dependent variable** (y) ip	pr	ir	er	cr
ei_gdp	0.0000	0.3996	0.7053	0.1043	0.0226
ei_pr	0.3328	0.0000	0.6311	0.3819	0.1676
ei_ir	0.0969	0.5407	0.7578	0.5451	0.0465
ip		0.6435	0.0236	0.0416	0.6191
pr	0.4581		0.0078	0.1597	0.4951
ir	0.1263	0.5380		0.0124	0.0042
er	0.0588	0.1059	0.0017		0.1972
cr	0.1786	0.1618	0.0114	0.0089	

H_0: x does not Granger cause y; H_1: x Granger causes y

Figure A.3: Impulse responses from Linear VAR, model with foreign variables, 95% confidence intervals

(a) Orthogonalized (b) Generalized

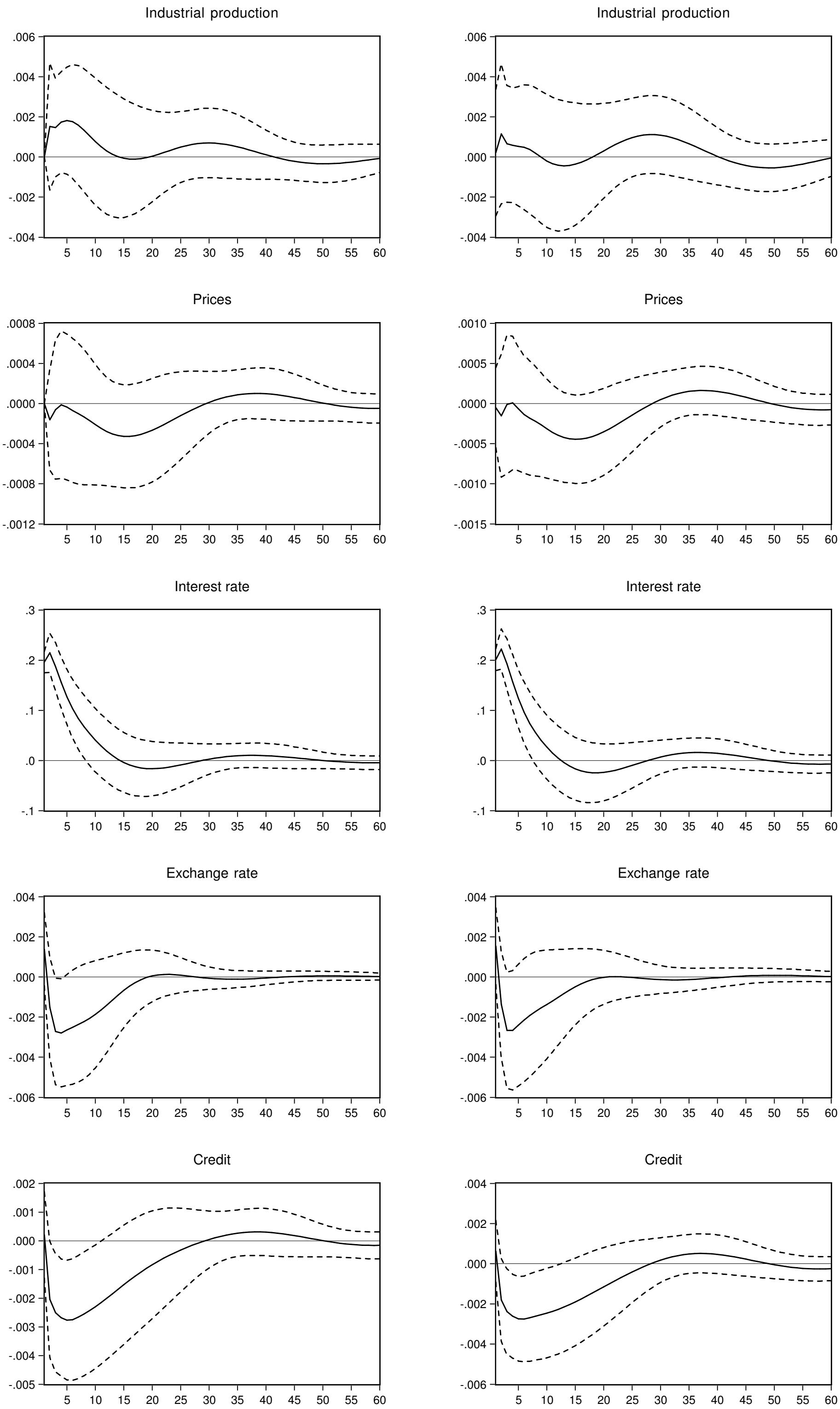


Figure A.4: Forecast Error Variance Decomposition

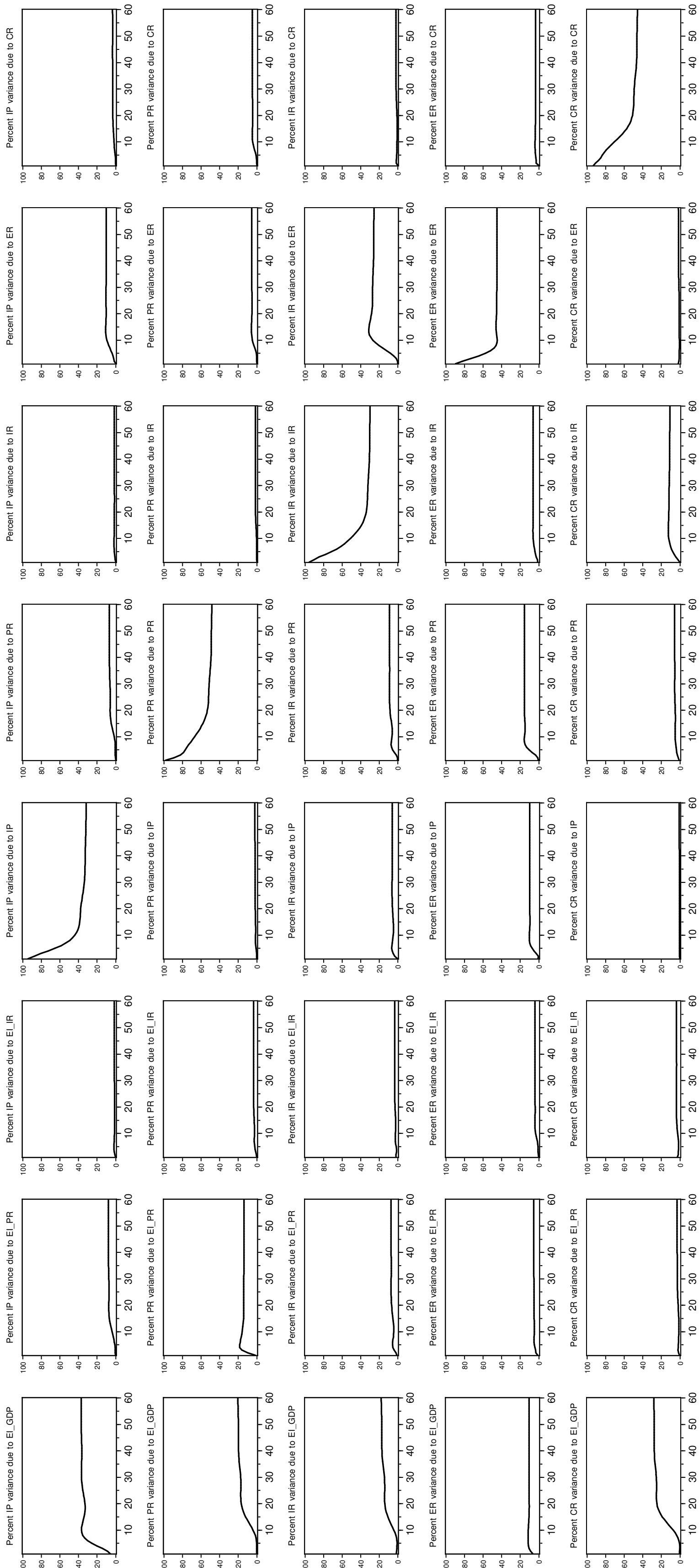


Figure A.5: Impulse responses from Linear VAR, model without foreign variables, 95% confidence intervals

(a) Orthogonalized (b) Generalized

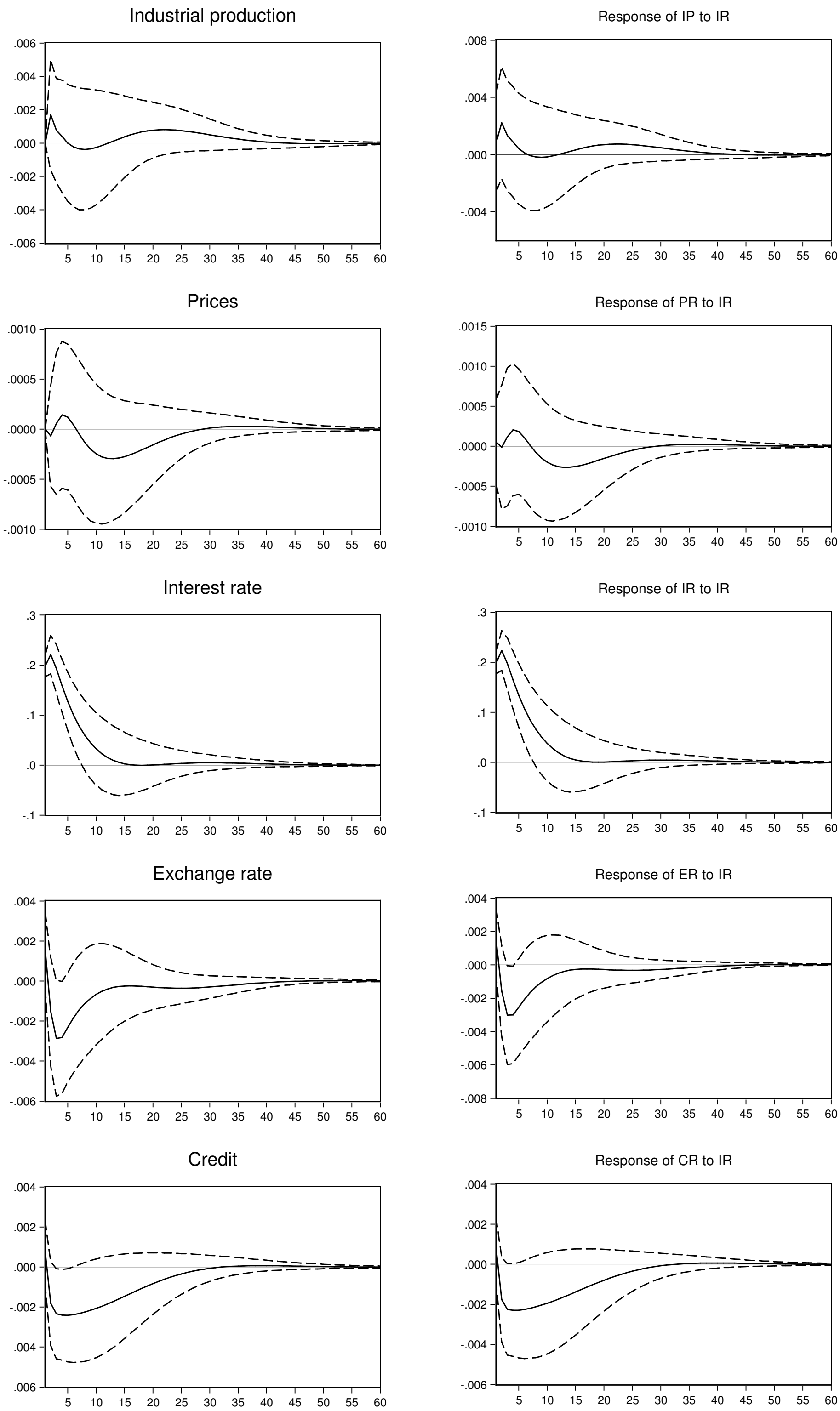


Figure A.6: Sensitivity analysis with GDP, high vs. low credit growth, 90% confidence intervals

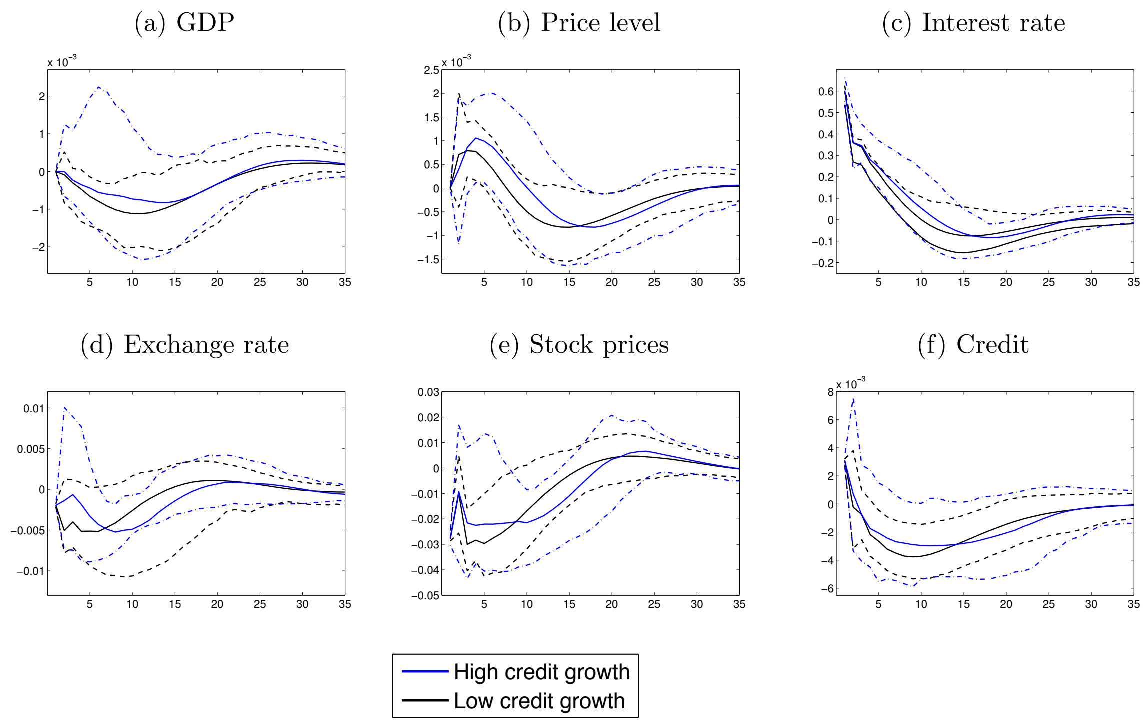


Figure A.7: Sensitivity analysis with GDP, positive vs. negative shock, 90% confidence intervals

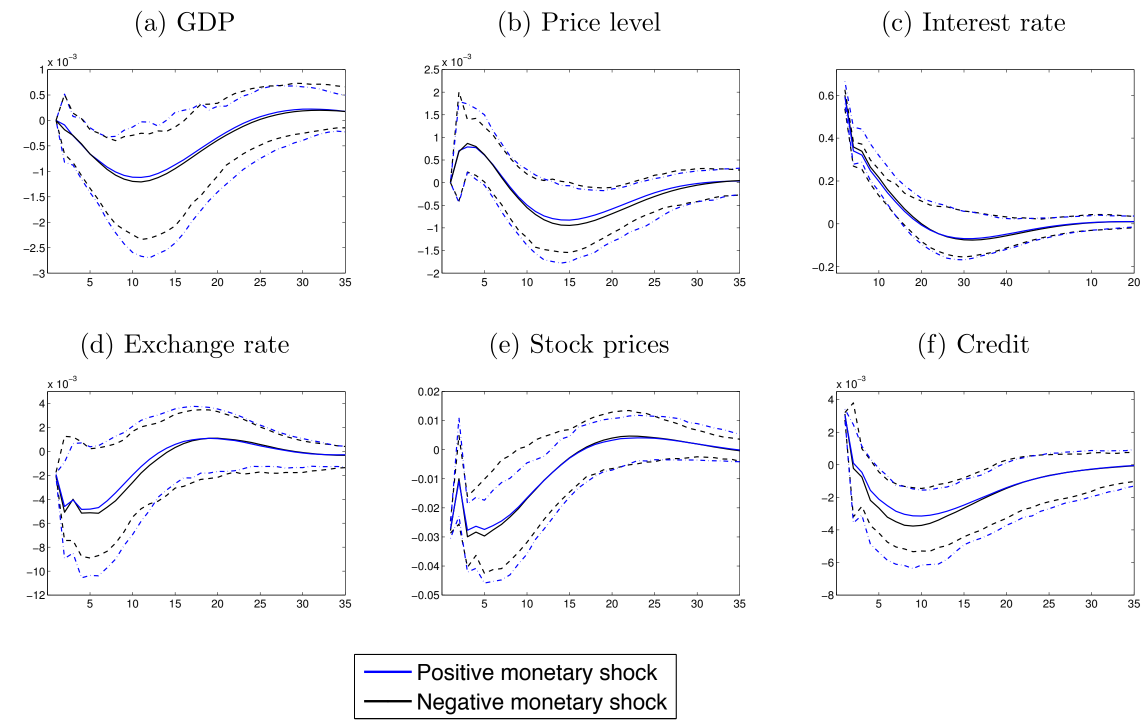
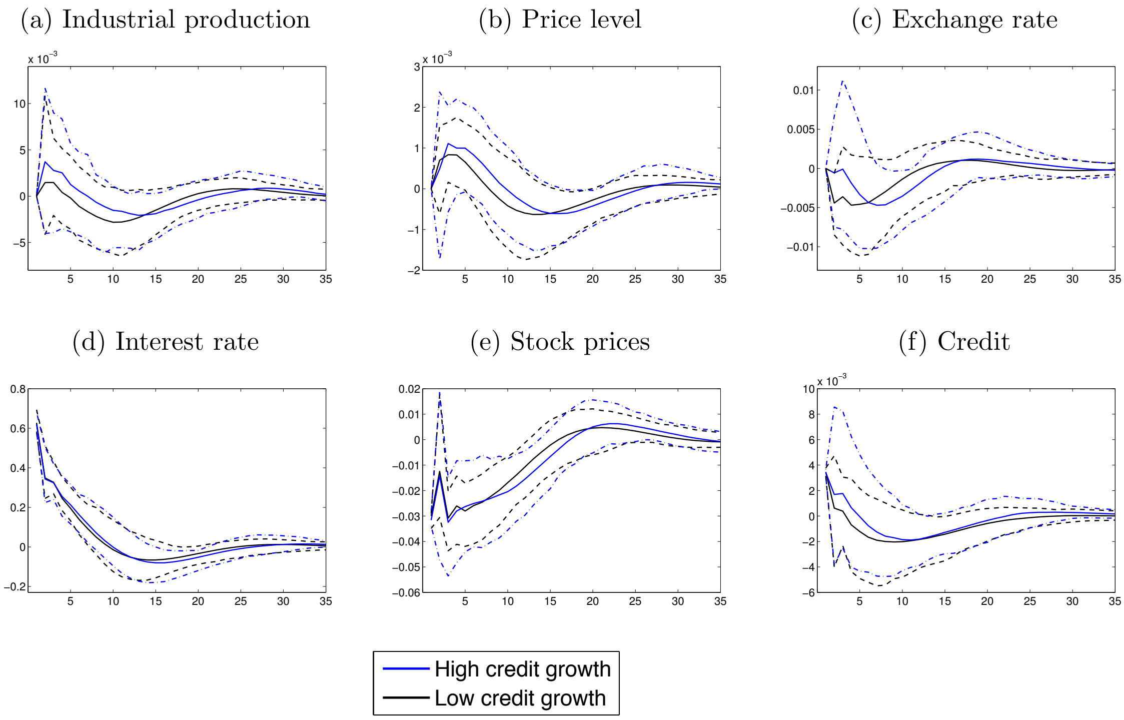


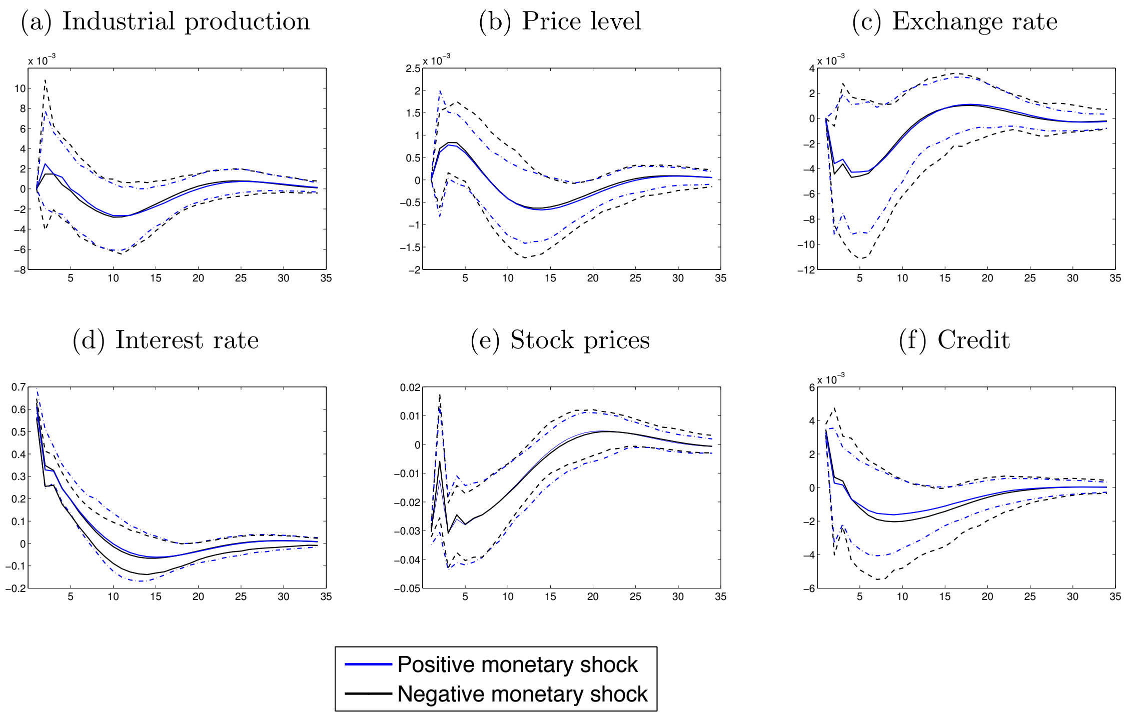
Figure A.8: Sensitivity analysis with the cycle length 60 months, high vs. low credit growth, 90% confidence intervals



(a) Industrial production (b) Price level (c) Exchange rate

(d) Interest rate (e) Stock prices (f) Credit

Figure A.9: Sensitivity analysis with the cycle length 60 months, positive vs. negative shock, 90% confidence intervals



(a) Industrial production (b) Price level (c) Exchange rate

(d) Interest rate (e) Stock prices (f) Credit

Appendix B

PCHVAR Data and Estimation Output

Table B.1: Dataset

Variable	Unit	Source
real GDP	national currency, 2000 prices	IMF IFS
price index	2005=100	IMF IFS
3 month money market rate	p.a.	IMF IFS
private credit/GDP	p.a.	World Bank
bank credit/bank deposits	p.a.	World Bank
bank overhead costs/total assets	p.a.	World Bank
net interest margin	p.a.	World Bank

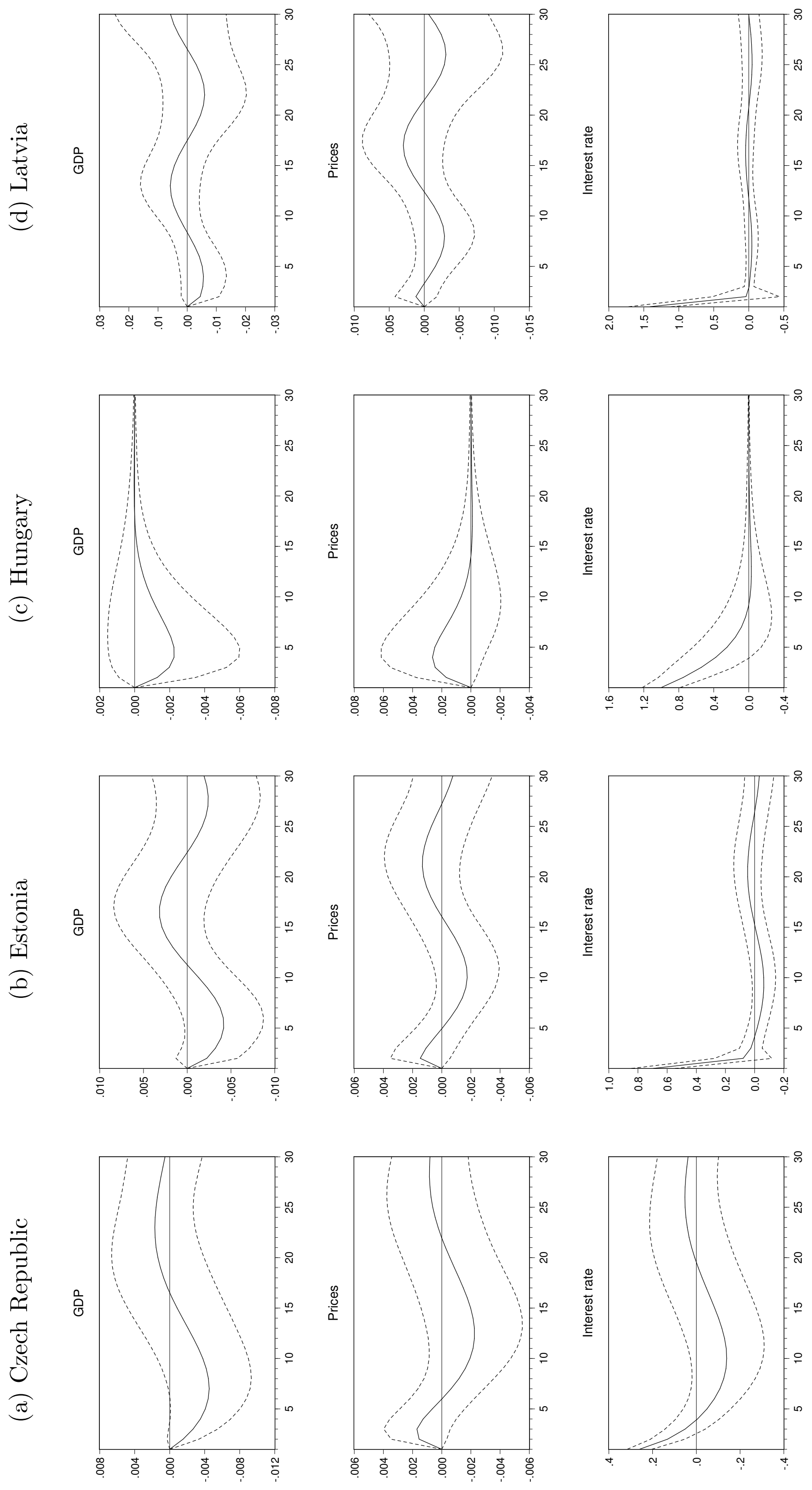


Figure B.1: Impulse responses obtained from linear VAR

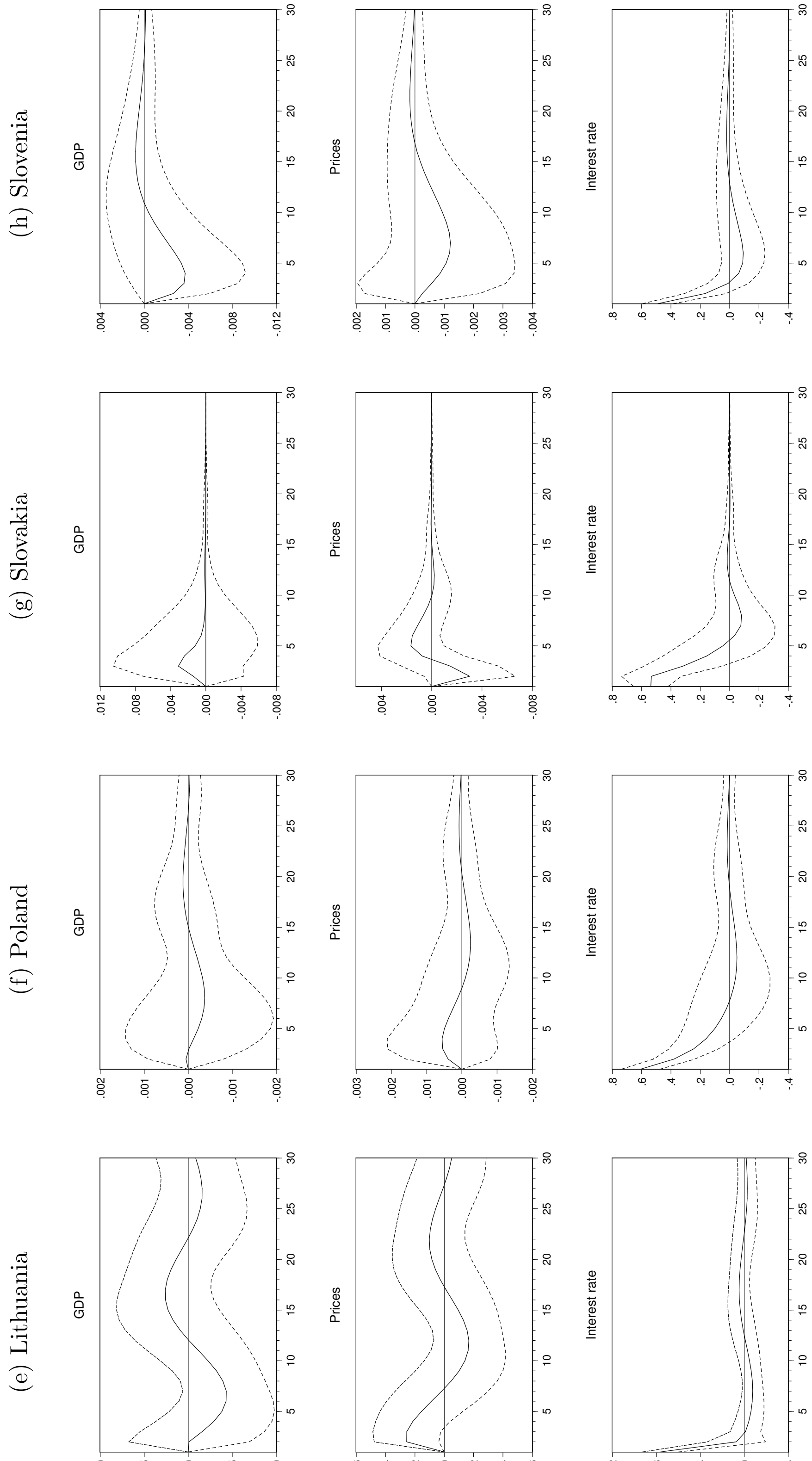
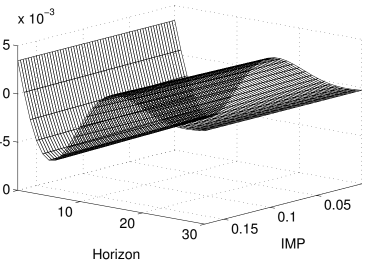


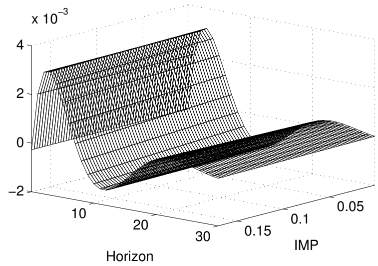
Figure B.1: Impulse responses obtained from linear VAR

Figure B.2: Impulse responses from PCHVAR conditioned on credit dependence, banking competition fixed at mean

(a) GDP



(b) Prices



(c) Interest rate

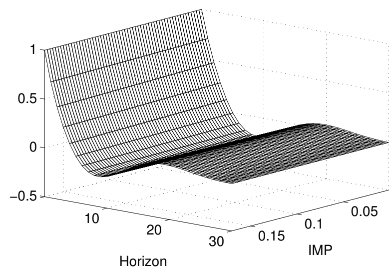
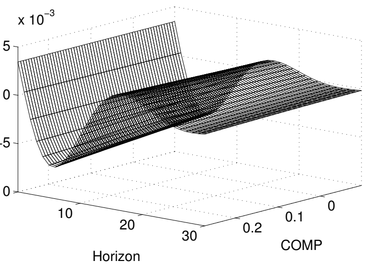
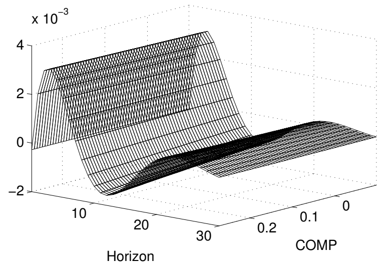


Figure B.3: Impulse responses from PCHVAR conditioned on banking competition, credit dependence fixed at mean

(a) GDP



(b) Prices



(c) Interest rate

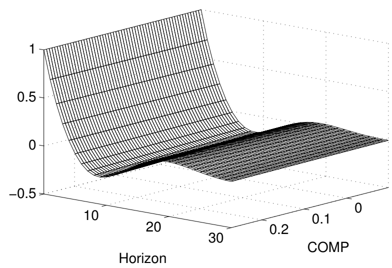
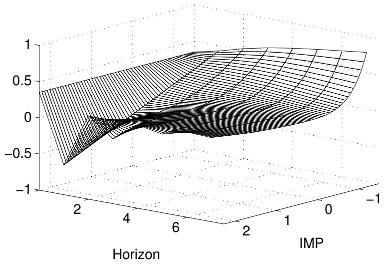
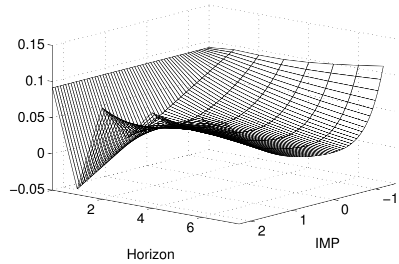


Figure B.4: Sensitivity analysis, impulse responses from PCHVAR conditioned on credit dependence

(a) GDP



(b) Prices



(c) Interest rate

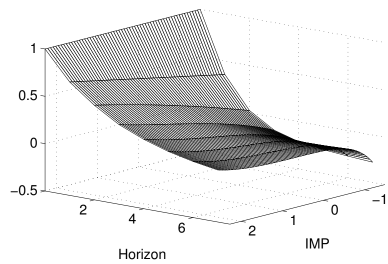
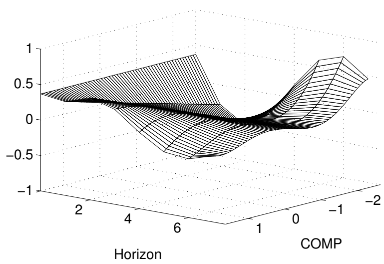
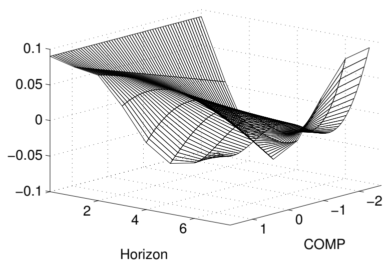


Figure B.5: Sensitivity analysis, impulse responses from PCHVAR conditioned on banking competition

(a) GDP



(b) Prices



(c) Interest rate

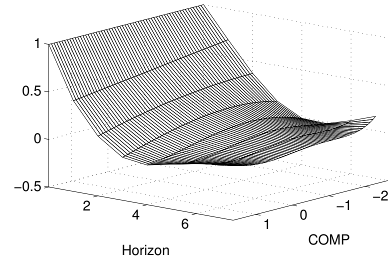
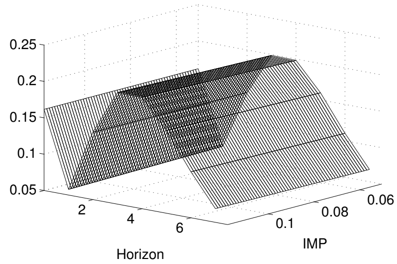
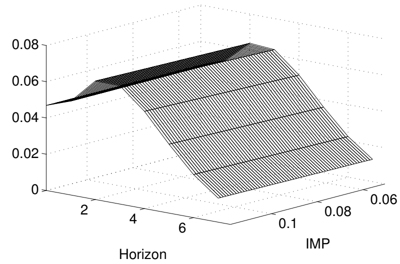


Figure B.6: Sensitivity analysis, impulse responses from PCHVAR conditioned on credit dependence, banking competition fixed at mean

(a) GDP



(b) Prices



(c) Interest rate

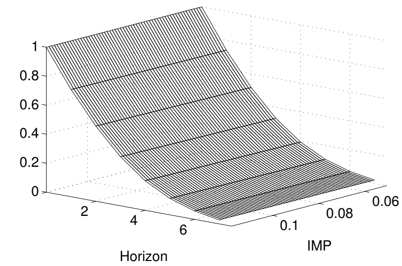
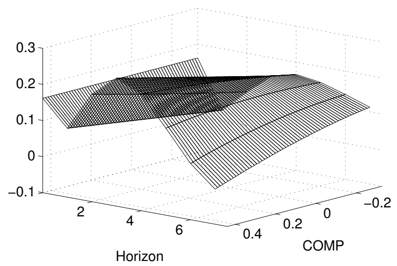
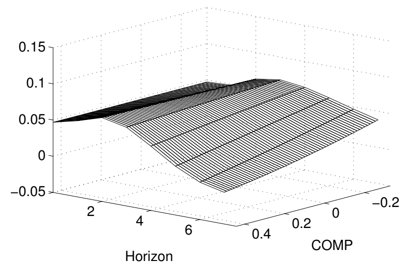


Figure B.7: Sensitivity analysis, impulse responses from PCHVAR conditioned on banking competition, credit dependence fixed at mean

(a) GDP



(b) Prices



(c) Interest rate